Understanding the Rise in Life Expectancy Inequality

Gordon B. Dahl, Claus Thustrup Kreiner,
Torben Heien Nielsen, Benjamin Ly Serena*

Final version

Abstract

We provide a novel decomposition of changing gaps in life expectancy between rich and poor into differential changes in age-specific mortality rates and differences in "survivability". Declining age-specific mortality rates increases life expectancy, but the gain is small if the likelihood of living to this age is small (ex-ante survivability) or if the expected remaining lifetime is short (ex-post survivability). Lower survivability of the poor explains half of the recent rise in inequality in the US and the entire rise in Denmark. Declines in cardiovascular mortality benefited rich and poor, but inequality increased because of differences in lifestyle-related survivability.

JEL Codes: H51, I14, J11

*Dahl: Department of Economics, University of California San Diego (gdahl@ucsd.edu); Kreiner: Center for Economic Behavior and Inequality, Department of Economics, University of Copenhagen (ctk@econ.ku.dk); Nielsen: Center for Economic Behavior and Inequality, Department of Economics, University of Copenhagen (thn@econ.ku.dk); Serena: Department of Economics, Copenhagen Business School (bse.eco@cbs.dk). We thank participants at the NBER workshop on *Income and Life Expectancy* in Boston, the workshop on *Health Inequalities* at the Copenhagen Business School and the workshop on *Behavioral Responses to Health Innovations and the Consequences for Socioeconomic Outcomes* at the University of Copenhagen for helpful discussions and comments. We are also grateful for discussions with Bo Honoré and Chris Ruhm and for comments by three anonymous referees and the editor, Ben Handel. Kristian Urup Olesen Larsen provided excellent research assistance. The Center for Economic Behavior and Inequality (CEBI) at the University of Copenhagen is supported by Danish National Research Foundation Grant DNRF134. This research was also supported by Novo Nordisk Foundation Grant NNF17OC0026542.

Life expectancy is strongly associated with socioeconomic status. This is a fundamental aspect of inequality in society and has important implications for the progressivity of public health and social security policies (Poterba, 2014; Auerbach et al., 2017). In many OECD countries, inequality in life expectancy has been rising.[1] Figure 1 displays estimates of life expectancy at age 40 across income tertiles for males and females in the United States and Denmark for 2001 and 2014.[2] The estimates are standard period life expectancies based on population-wide register data on mortality and income (see Chetty et al. 2016). The gap in life expectancy between rich (top tertile) and poor (bottom tertile) males is around 8 years in both the US and Denmark in 2001. Over the short period from 2001 to 2014, this inequality increased by 1.7 years in the US and 0.9 years in Denmark. The gap between rich and poor females stayed constant in Denmark over this period, but also increased by about 1.8 years in the US.

[FIGURE 1 HERE]

The driving forces behind the recent trends in life expectancy inequality remain unclear. Several studies focus on comparing changes in age-specific mortality rates by socioeconomic status,[3] but how do these changes translate into trends in life expectancy inequality? Will a larger drop in the mortality rates of the poor than the rich necessarily reduce the gap in life expectancy?

We provide a novel decomposition which links changes in life expectancy inequality to underlying changes in mortality inequality. This decomposition splits the rise in life expectancy inequality into differential mortality trends between rich and poor and a common mortality trend.[4] The common mortality trend affects inequality in life expectancy because of differences in "survivability" of the rich and the poor. Survivability is a factor that translates changes in mortality into changes in life expectancy. It is computed from initial age-specific mortality rates and measures

[1] See Waldron (2007); Case and Deaton (2015); Chetty et al. (2016); Currie and Schwandt (2016); Auerbach et al. (2017); Bor et al. (2017); Hederos et al. (2017); Kreiner et al. (2018); Kinge et al. (2019).

[2] We focus on life expectancy by income at age 40 as in other recent work (e.g., Chetty et al., 2016; Kreiner et al., 2018; Kinge et al., 2019). Other research looks at life expectancy at birth; to do this, researchers have used county-level income because individual income at birth is an inadequate proxy for social class. This approach includes changes in mortality inequality at younger ages, which in isolation have reduced life expectancy in recent decades in some countries (Currie and Schwandt, 2016; Baker et al., 2019). We discuss sensitivity analyses that include changes in mortality throughout the life span in Section 7.

[3] See Lleras-Muney (2005); Snyder and Evans (2006); Cutler et al. (2011); van den Berg et al. (2017); Mackenbach et al. (2018); Baker et al. (2019); Montez et al. (2019); Attanasio and Nielsen (2020).

[4] Many decomposition methods and applications exist, but as we describe in Section 2, none have decomposed life expectancy between groups and over time into these two components before.

the likelihood of surviving until a given age multiplied by the expected remaining life years after surviving this age. Intuitively, a person only benefits from a reduction in an age-specific mortality rate if they have survived until this age (ex ante effect) and, if so, the benefit is the expected extra life years thereafter (ex post effect).

Increasing life expectancy inequality can arise because the gap in mortality rates between the rich and the poor increases, for example if medical advances differentially benefit the rich or if health behaviors differentially worsen for the poor (Cutler et al., 2006; Jayachandran et al., 2010; Cutler et al., 2011; Case and Deaton, 2015; Moscelli et al., 2018). But it can also arise because the rich have higher survivability due to lower initial mortality rates. As we show, this implies somewhat paradoxically that the gap in life expectancy between the rich and poor can increase, even if the gap in mortality rates is constant or declining. Thus, both differential mortality rate changes across groups and existing mortality inequality across groups (survivability differences) are key determinants of changes in life expectancy inequality over time.

Empirically, we find that it is important to account for survivability when evaluating changes in life expectancy inequality in both the US and Denmark. In the US, half of the rise in inequality for forty-year old males shown in Figure 1 is due to larger reductions in mortality rates for the rich than the poor, while the other half is due to differences in their survivability. For Danish males in Figure 1, life expectancy inequality increased, even though mortality rates have fallen more for the poor. The explanation for this apparent puzzle is that survivability strongly favored the rich, more than offsetting the effect of differential mortality rate changes. For females in both countries, survivability plays a similarly important role.

Motivated by influential work on cause-specific mortality (Case and Deaton, 2015; Cutler and Kadiyala, 2003), we next explore how this interacts with cause-specific survivability. We extend our decomposition of life expectancy inequality into four broad death categories: cardiovascular, cancer, miscellaneous lifestyle (smoking, alcohol, and obesity related deaths not classified as cardiovascular or cancer), and other causes.[5] In Denmark, where we are able to link cause-specific deaths to income, we observe a large drop in cardiovascular mortality rates, and more so for the poor than the rich. Despite the larger reduction in mortality of the poor, the gap in life expectancy

[5] While this is a natural grouping of the major causes of death, we note that cardiovascular and cancer deaths are often related to lifestyle (e.g., lung cancer).

increased because the poor were more likely to die of miscellaneous lifestyle-related causes before benefiting from the reduction in cardiovascular mortality (ex ante survivability), and because those surviving to benefit gained fewer extra life years as a result of higher lifestyle-related mortality in subsequent ages (ex post survivability). Defining socioeconomic class by education level instead of income allows us to perform an analogous exercise for the US, where a similar conclusion emerges.

Our decomposition demonstrates the value of linking the change in life expectancy inequality to the underlying change in age-specific mortality rates. Trends in age-specific mortality rates of the rich and poor are informative about changes in underlying health status, while trends in life expectancy are a relevant measure of the associated welfare effects. Looking at each of these in isolation misses an important link – survivability – and, as demonstrated by our empirical results, can lead to misleading conclusions. Moreover, our cause-specific findings make clear that the driving forces for changes in mortality inequality can be very different from those for changes in life expectancy inequality. At the end of the paper, we discuss several additional measurement issues which could affect interpretation.

From a policy perspective, the lesson is that mortality rate trends do not necessarily provide an accurate indication of how to best combat inequality in life expectancy. Investments in public health, such as new medical technologies, can improve mortality rates and generate overall improvements in life expectancy. However, a fundamental tension exists, because mortality rate improvements can result in a rise in life expectancy inequality even when they favor the poor. Thus, health policies that reduce mortality inequality can paradoxically lead to more inequality in longevity and make social security systems and other age-related transfer payments more regressive.

1 An Illustrative Empirical Example

To illustrate how changes in life expectancy inequality are affected by mortality changes and survivability, consider the impact of the observed change in the one-year mortality rates at age 60 from 2001 to 2014 on life expectancy of males at age 40. Column 1 of Table 1 shows these changes for the rich and poor in the US and Denmark. While the decrease in mortality at age 60 is the same across the income groups within each country, the impact on life expectancy is not. The reason is survivability. In the US, 95% of 40-year-old males are predicted to survive to age 60 if they are rich, compared to 84% of the poor (column 2). Hence, poor individuals are less likely than rich individuals to survive long enough to benefit from the reduction in mortality which occurs at age 60 (ex ante survivability). For those who do survive past age 60, remaining life expectancy is 23.5 years for the rich versus 19.0 for the poor (column 3). In other words, the poor are more likely to die sooner if they survive to age 61 compared to the rich (ex post survivability).

[Table 1 HERE]

By multiplying the two survivability components, we obtain the total survivability effect in column 4. This shows that the rich would gain 22.3 years if the mortality rate went from 1 to 0 at age 60, while the poor would only gain 16.0 years. Finally, by multiplying the observed changes in mortality (column 1) by survivability (column 4), we obtain the change in life expectancy in column 5. This calculation reveals that even with the same change in mortality (-.002) for the rich and the poor, the rich gain 1.4 times more years of life expectancy (.045 versus .032 years). A similar pattern holds for Denmark. It further follows from the example (by continuity) that larger declines in mortality of the poor than the rich, which narrows the gap in mortality, may expand gaps in life expectancy. Indeed, this is what happens in practice for Denmark, when we consider all changes in age-specific mortality rates in Section 4.

The example in Table 1 illustrates how mortality rate changes at a given age map into changes in life expectancy and the role played by survivability. The next section shows this more generally with mathematical formulas that are later used to assess the empirical importance of survivability.

2 Decomposition Formulas

A standard formula for life expectancy measured at age $\underline{a}$ is

$$\begin{aligned} LE_{\underline{a}} &= \underline{a} + (1 - M_{\underline{a}}) + (1 - M_{\underline{a}})(1 - M_{\underline{a}+1}) + ... \\ &= \underline{a} + \sum_{a=\underline{a}}^{\infty} \prod_{i=\underline{a}}^{a} (1 - M_i) \end{aligned} \tag{1}$$

where M_a is the mortality rate at age a. This equation and the following decomposition formulas apply both to period life expectancy and cohort life expectancy, the difference being how age-specific mortality rates are estimated. Cohort life expectancy uses mortality of the same cohorts over time, while period life expectancy uses mortality of different cohorts in a given period. In our empirical application, we focus on period life expectancy in line with recent studies of inequality over time (Chetty et al., 2016; Currie and Schwandt, 2016; Hederos et al., 2017; Kreiner et al., 2018; Kinge et al., 2019).[6]

Differentiating equation (1) with respect to mortality rates and summing over all ages yields the following first-order approximation for the change in life expectancy (see Appendix A.3):

$$\Delta LE_{\underline{a}} \approx -\sum_{a=\underline{a}}^{\infty} \Delta M_a \cdot X_a \text{ where } X_a \equiv S_a \cdot R_a \tag{2}$$

where $S_a \equiv \prod_{i=\underline{a}}^{a-1} (1 - M_i)$ is the probability of survival from age $\underline{a}$ to age a and $R_a \equiv 1 + \sum_{j=a+1}^{\infty} \prod_{i=a+1}^{j} (1 - M_i)$ is remaining life expectancy after surviving age a, in which case the individual lives one extra year for sure and from age a+1 can expect to live the additional years given by the last term in the definition.

Equation (2) expresses the change in life expectancy as the product of changes in age-specific mortality rates (ΔM_a) and survivability (X_a), where survivability is a summary measure of initial mortality rates that includes the ex ante (S_a) and ex post (R_a) terms. This aligns with the intuition from Table 1, the only difference being that the formula sums the changes over all possible ages. Equation (2) is a mathematical identity for infinitesimal mortality changes and a first-order approximation for actual changes.[7]

By applying equation (2) for rich and poor and using a little algebra (see Appendix A.3), we obtain the following decomposition formula for the change in life expectancy inequality between the rich (superscript r) and the poor (superscript p):

$$\Delta LE_{\underline{a}}^r - \Delta LE_{\underline{a}}^p \approx \sum_{a=\underline{a}}^{\infty} \overline{X_a} \cdot (\Delta M_a^p - \Delta M_a^r) + \sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot (X_a^p - X_a^r) \tag{3}$$

[6] Since period life expectancy is a summary measure of cross-sectional mortality rates at a given point in time, it is often used to study trends in inequality. In a steady state with constant age-specific mortality rates, period life expectancy would equal the observed average life length. Therefore, comparing period life expectancy at two points in time, as done in the literature, is effectively comparing expected longevity between two (artificial) steady states and does not, for example, capture the full benefits on actual longevity from health improvements that take place in this time span. For a further discussion of period life expectancy and cohort life expectancy see Guillot (2011).

[7] The age decomposition in (2) is a simple first order approximation, which works well in our empirical application (see Appendix Figure A.2). One could alternatively use an Arriaga age decomposition (Arriaga, 1984), which can be rewritten as equation (2) plus an interaction term $\Delta R_a \cdot \Delta M_a \cdot S_a$. The extra term reduces approximation error, but cannot be attributed to any particular age and hence is difficult to interpret. Our key decomposition formula (3) is based on the age decomposition in equation (2). The main empirical results are unchanged if we include the interaction term in equation (2).

where $\overline{X_a}$ is the average survivability of the rich and the poor, while $\overline{\Delta M_a}$ is the average change in their mortality rates. The formula decomposes the change in life expectancy inequality into two terms. The first term isolates the contribution from differential changes in the mortality rates of rich and poor, holding survivability fixed at the average across income groups ($X_a^r = X_a^p = \overline{X_a}$). The second term isolates the contribution from differences in survivability between the rich and poor, holding mortality rate changes fixed across the two groups ($\Delta M_a^r = \Delta M_a^p = \overline{\Delta M_a}$).

In our empirical example from the US and Denmark in Section 1, the observed changes in mortality rates were the same for the rich and poor (the first term in the formula is zero). In this special case with a common mortality trend, the change in life expectancy inequality is driven entirely by differences in survivability. In the general case, however, differential mortality trends may also contribute to changes in life expectancy inequality (through the first term in the formula).

Our decomposition is related to the seminal work of Kitagawa (1955), which has been applied and extended in various forms.[8] Normally, the purpose is to decompose the difference in crude rates (e.g., crude death rates) between two populations into differences in the composition of characteristics in the population and differences in characteristic-specific rates. Our formula links changes in life expectancy inequality to differential mortality trends between rich and poor, and the overall mortality trend, which affects life expectancy inequality through differences in survivability. To our knowledge, nobody has decomposed life expectancy inequality between groups and over time in this way before.[9]

The Oaxaca-Blinder decomposition widely used in economics is a related idea, and decomposes the difference in outcomes between two groups into differences in mean characteristics across the groups and the differential effect of characteristics across groups (Oaxaca, 1973; Blinder, 1973). Technically, our decomposition differs from the standard Oaxaca-Blinder decomposition, in part because it sums over a variety of ages, and in part because it uses the combined means for both groups ($\overline{X_a}$ and $\overline{\Delta M_a}$, rather than separate means by group). We further take advantage of the

[8] See Fortin et al. (2011) and Canudas Romo (2003) for discussions of decomposition methods in economics and demography, respectively.

[9] Jdanov et al. (2017) decompose current life expectancy inequality levels into differences in historical mortality trends and initial differences in mortality. However, they do not consider that the impact of age-specific mortality trends on life expectancy depends on pre-existing differences in mortality at other ages (survivability), which is the main point of our decomposition.

fact that with two equally-sized groups, we can use the combined means for both groups without needing to account for any covariance terms in the decomposition formula.

Using the definition $X_a \equiv S_a \cdot R_a$, we can further decompose the last term in equation (3) into the contributions from ex ante and ex post survivability, and arrive at the following decomposition (see Appendix A.3):

$$\Delta LE^r_{\underline{a}} - \Delta LE^p_{\underline{a}} \approx \underbrace{\sum_{a=\underline{a}}^{\infty} \overline{X_a} \cdot (\Delta M^p_a - \Delta M^r_a)}_{\Delta\text{Mortality}} + \underbrace{\sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{R_a} \cdot (S^p_a - S^r_a)}_{\text{Ex ante survivability}} + \underbrace{\sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{S_a} \cdot (R^p_a - R^r_a)}_{\text{Ex post survivability}} \quad (4)$$

where $\overline{R_a}$ and $\overline{S_a}$ are the average ex ante and ex post survivability of the rich and the poor.

3 Data

Our empirical analysis is based on mortality data for the US and Denmark from 2001 to 2014. One benefit of using data from both the US and a European country is that they differ in the amount of income inequality and in the private versus public provision of health care. US data by income class and age comes from the study by Chetty et al. (2016), which uses the universe of IRS tax returns. The Danish data takes advantage of population-wide administrative registers collected by Statistics Denmark. We measure income two years prior to age-specific mortality, using the sum of household gross income minus Social Security and disability benefits.[10]

The Danish data has several advantages over the US data for our study. First, the Danish data covers the entire population, while the US data does not include individuals with zero earnings. Second, with the detailed Danish data it is not necessary to impute mortality rates by socioeconomic status after retirement (which the US analysis does using Gompertz approximations). The third, and most important, advantage is that the Danish data can link cause of death to income data. This facilitates our cause-specific decomposition of changes in life expectancy inequality. Additional information about the datasets are provided in Appendix A.1.

[10]Recent work suggests it is important to account for income mobility over the life cycle when computing inequality in life expectancy (Kreiner et al., 2018). In the Appendix we find the change in life expectancy inequality is smaller if we account for income mobility, but the relative importance of survivability is largely unchanged.

4 Empirical Importance of Differences in Survivability

The simple empirical example for the US and Denmark provided in Table 1 illustrates the role of survivability for rising life expectancy inequality. For both countries, the change in mortality at age 60 is identical for the rich and the poor. However, this is not true at all ages. The left panel of Figure 2 plots the change in mortality from 2001-2014 over 5-year age bins for the rich and the poor. An interesting contrast emerges between the two countries. In the US, starting around age 60, mortality rates fall more dramatically for the rich than the poor for both males and females. In Denmark, the drop in age-specific mortality rates has, for most ages, actually favored the poor.

[FIGURE 2 HERE]

The right panel plots survivability, which is the product of survival until a given age (ex ante survivability) multiplied by remaining life expectancy beyond that age (ex post survivability). Survivability always favors the rich, both in the US and Denmark and for both males and females. While not shown in the figure, both ex ante and ex post survivability favor the rich at every age. In other words, the rich are more likely to be alive to benefit from a drop in mortality at any age, and have more remaining years to benefit at any age. The gap in survivability is largest for males and narrows with age, but it never completely disappears. This means that a common drop in mortality rates at any age will mechanically favor the rich, widening inequality in life expectancy.

To assess the contributions of each component more systematically, we use the decomposition formulas (3) and (4). Focusing first on the US, we decompose the change in inequality in life expectancy between the rich and poor from 2001 to 2014. As can be seen in Figure 3, for males, the gap in life expectancy between the rich and poor increased in total by approximately 1.4 years. Roughly half of this increase is attributable to larger drops in mortality for the rich than the poor. The other half is due to survivability which favors the rich, with ex ante and ex post survivability playing equal roles.[11]

[11] One way to benchmark the contribution of ex ante versus ex post survivability is to calculate the effect survivability would have if the overall decline in mortality across all ages instead had been concentrated at a single, specific age. For US males, if the entire mortality decline was concentrated at age 40, there would be a 1.5 year increase in life expectancy inequality between the rich and poor. All of the effect would operate through ex post survivability since there are no differences in ex ante survivability at age 40. If the entire mortality decline was instead concentrated at age 85, there would have been a 0.4 year increase in life expectancy inequality, with almost all of the gap being driven by ex ante survivability. These benchmarks compare to the actual survivability effect of a 0.7 year increase, and indicate that a general reduction in mortality has a larger effect if it occurs at younger ages.

[FIGURE 3 HERE]

For females in the US, there is a 1.5-year increase in life expectancy inequality. Over two-thirds of this rise is attributable to drops in mortality favoring the rich. The remaining is mostly due to ex ante survivability, i.e., a larger fraction of poor females not living long enough to benefit from reductions in mortality.

A different pattern emerges for Denmark in Figure 3, where age-specific improvements in mortality have favored the poor. For males, changes in mortality reduce inequality by a sizeable 0.8 years. In spite of this, overall life expectancy inequality increases by 0.7 years. This reversal in sign occurs because survivability strongly favors the rich. Ex ante survivability accounts for a 0.9-year increase in life expectancy inequality, and ex post survivability accounts for another 0.6-year increase. So while drops in age-specific mortality rates favored the poor, the poor did not live long enough to benefit and did not gain as many years of remaining life compared to the rich.

For females in Denmark, there is virtually no change in life expectancy inequality. This is despite the fact that drops in mortality contributed almost a full year more to life expectancy for the poor than the rich during this 14-year period. But offsetting the drop in mortality favoring the poor was an equally large contribution from survivabililty favoring the rich. Ex ante and ex post survivability each account for roughly a half-year increase in life expectancy inequality.

5 Contributions from Cause-Specific Mortality

The previous section highlights the importance of survivability for understanding changes in life expectancy inequality. We now explore how recent changes in cause-specific mortality have interacted with cause-specific survivability to affect life expectancy inequality. To do this, we extend our decomposition approach so that the components of equation (4) are cause specific, where the subscript c denotes the cause (see Appendix A.5):

$$\Delta LE^r_{\underline{a}} - \Delta LE^p_{\underline{a}} \approx \underbrace{\sum_{a=\underline{a}}^{\infty} \overline{X}_a \cdot \sum_c \left(\Delta M^p_{a,c} - \Delta M^r_{a,c}\right)}_{\Delta\text{Mortality}} + \underbrace{\sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{R_a} \cdot \sum_c (S^p_{a,c} - S^r_{a,c})}_{\text{Ex ante survivability}} + \underbrace{\sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{S_a} \cdot \sum_c (R^p_{a,c} - R^r_{a,c})}_{\text{Ex post survivability}} \quad (5)$$

where $\Delta M^r_{a,c}$ and $\Delta M^p_{a,c}$ are cause-specific mortality rates of the rich and poor, $S^r_{a,c}$ and $S^p_{a,c}$ are their probabilities of not dying of cause c before age a, and $R^p_{a,c} - R^r_{a,c}$ denotes the contribution of cause c to the difference in remaining life expectancy between the rich and poor at age a.

This cause-specific decomposition of changes in life expectancy inequality is novel. Note that it deviates from an Arriaga decomposition, which decomposes life expectancy by cause-specific mortality differences (Kinge et al., 2019; Ho and Hendi, 2018) either between groups or over time, and which attributes age-specific contributions based on cause-specific mortality differences at that age. Our decomposition allows us to study differences in life expectancy between groups and over time, and captures that the impact of cause-specific changes in mortality rates at one age depend on initial differences in cause-specific mortality at other ages (cause-specific survivability).

The key assumption needed for this type of decomposition is independence across cause-specific changes in mortality rates. This simplifying assumption, which is generally made in cause-specific decompositions of life expectancy and in related work on competing risks, implies that improvements in cardiovascular mortality rates do not, for example, affect changes in cancer mortality rates. One exception is Honoré and Lleras-Muney (2006), which provides bounds in a competing risk model without assuming independence. They estimate changes in cancer and cardiovascular mortality over time without assuming independence, and find that the improvements in cancer are larger compared to estimates which assume independence.

We decompose changes in life expectancy inequality into four broad categories: cardiovascular, cancer, miscellaneous lifestyle (smoking, alcohol, and obesity related deaths not classified as cardiovascular or cancer), and other causes.[12] The separation into these four categories is motivated by cardiovascular disease and cancer being leading causes of death, and because what we label as "miscellaneous lifestyle" has been the focus of influential research and the debate on deaths of despair (Fuchs, 1974; Cutler et al., 2011; Case and Deaton, 2017; Ruhm, 2018; Haan et al., 2019). When interpreting our results, we note that many deaths due to cardiovascular disease and cancer could also be lifestyle related (e.g., lung cancer).

[FIGURE 4 HERE]

[12] Miscellaneous smoking/alcohol/obesity-related diseases consist of COPD (ICD-10: J40-J44), alcohol abuse (F10), alcoholic liver disease and cirrhosis (K70, K74), alcohol-induced chronic pancreatitis (K86.0), diabetes (E10-E14), obesity (E66).

The top panels of Figure 4 shows time trends in these four cause-specific mortality rates for Denmark, separately for males and females. The dramatic reduction in cardiovascular mortality observed for both genders mirrors a global trend, which has been associated with medical advances in the treatment and prevention of cardiovascular disease (Deaton, 2013; Cutler and Kadiyala, 2003; Likosky et al., 2018). While mortality rates for cancer, lifestyle, and other causes decline modestly over our sample period, these drops are overshadowed by the halving of cardiovascular mortality.

The bottom panels of Figure 4 plot the cause-specific decomposition results for Denmark based on income groups. The large bars in panels C and D replicate the decomposition already presented for Danish males and females in Figure 3. The smaller bars further decompose each of the large bars into contributions from the different causes of death. Similar data on cause-specific mortality by income is not readily available for the US, but in the next section we discuss cause-specific results when using education as a proxy for social status, which is possible to do for both countries.

We first look at the importance of cause-specific mortality rate changes for rich versus poor Danish males. The leftmost small bar in panel C shows that the big drop in cardiovascular mortality contributed to a large decline of 0.7 years in the rich-poor life expectancy gap. This is the main reason why changes in mortality rates reduced inequality in Denmark. Changes in miscellaneous lifestyle-related death causes also reduced inequality, while trends in cancer mortality and other death causes both contributed little to changes in life expectancy inequality.

If the decline in cardiovascular mortality favored the poor, what caused overall inequality to increase? The answer is differences in survivability due to other causes of death, and in particular miscellaneous lifestyle causes of death. The poor were more likely to die of death causes related to smoking, alcohol, and obesity before benefiting from the specific reduction in cardiovascular mortality (ex ante survivability), and those surviving to benefit had fewer remaining life years because of higher mortality from lifestyle-related causes of death in subsequent ages (ex post survivability). Cardiovascular mortality also contributes to the survivability gap between the rich and poor, while there is little impact from cancer.

The third set of cause-specific bars in panel C shows the total change in life expectancy inequality between the rich and poor. Miscellaneous lifestyle and other causes of death are both large contributors to the rising gap. This is not because of increasing gaps in lifestyle mortality rates, which actually went the other direction, but because initial differences in survivability from lifestyle mortality meant that the overall drop in cardiovascular mortality had a smaller impact on life expectancy for the poor than the rich. A broadly similar explanation holds for other death causes as well. It is also worth noting that cancer mortality had small effects, both in terms of mortality changes and survivability, even though it is the second leading cause of death.

Turning to females in panel D of Figure 4, a similar pattern emerges. While cardiovascular mortality dropped more for the poor than the rich (the leftmost small bar for females), higher initial mortality and thus lower survivability for the poor offsets this favorable mortality development. Lifestyle causes account for more than one third of the contribution from differences in survivability, with cardiovascular and other causes accounting for most of the remaining.

These results show that widening inequality in Denmark is not happening because lifestyle-related mortality differentially worsened for the poor, but because reductions in cardiovascular mortality made existing differences in lifestyle-related survivability more consequential. The cause-specific decompositions illustrate another important point. Causes of death that contribute most to changes in mortality inequality are not necessarily the most important contributors to changes in life expectancy inequality.

6 Education and Gender Based Decompositions

Following recent work on measuring inequality in life expectancy, we proxy social status by position in the income distribution (Chetty et al., 2016; Kreiner et al., 2018; Kinge et al., 2019). The same general conclusions apply if we instead use educational attainment, see Appendix Section A.6. While we cannot link cause of death to income data for the US, we can link cause of death to education level in both the US and Denmark. For both countries and both sexes, the drop in mortality rates related to cardiovascular disease helped reduce the gap in life expectancy between the low-educated and high-educated. Nevertheless, total inequality in life expectancy increased. A key reason is the difference in survivability between the low and high-educated in both countries, where differences in lifestyle-related causes of death play a major role.

Using education, we may also assess the external validity of our findings. In Appendix Section A.7, we provide a decomposition of the trend in life expectancy inequality for nine Western European countries where education data is available. The decomposition results for these countries are broadly similar to the findings for Denmark, suggesting that the results for Denmark are representative of Western European countries.

Turning to gender, we ask how the female-male gap in life expectancy has evolved over this same time period. Females have longer life expectancies than males, but the gap has decreased in recent years (Goldin and Lleras-Muney, 2019). In Appendix Figure A.5, we use our decomposition formula (3), but applied to females and males instead of rich and poor, to assess the importance of survivability. Changes in mortality rates declined faster for males versus females between 2001 and 2014, with almost a full-year decline in the US and a 1.2-year decline in Denmark. But counteracting this drop was lower survivability for males. In other words, lower ex ante and ex post survivability meant that males did not live long enough to benefit from the differential change in mortality and gained fewer life years upon surviving a given age.

7 Measurement Issues

Our results highlight the importance of survivability in linking changes in mortality and life expectancy inequality. There are several additional measurement issues in the linking of these outcomes which we summarize here, with details provided in Appendices A.8-A.12.

We focus on absolute changes in mortality rates as is mostly done in the literature (Currie and Schwandt, 2016; Case and Deaton, 2015; Kinge et al., 2019). This allows our insight to be directly applied to that work, and also has the advantage that the decomposition in levels is additively separable into easily interpretable components. One might ask whether differential relative changes in mortality rates between rich and poor would map one-to-one into differential changes in their life expectancy. This turns out not to be the case. As we show, differences in survivability are still important and life expectancy may rise more for the rich than the poor, both in absolute terms and in relative terms, even when relative changes in mortality rates are the same for the two groups.

Life expectancy estimates do not account for quality of life. A quantitative approach to address this is to compute disability-adjusted life expectancy, which puts lower weight on life years with a high expected burden of disability (Marmot et al., 2010). When doing this, based on World Health Organization (WHO) estimates of years lived with disability, the rise in inequality becomes smaller. More importantly, the relative role played by differences in survivability between rich and poor in explaining the rise in life expectancy inequality is unchanged.

We follow previous literature (Chetty et al., 2016; Kreiner et al., 2018; Kinge et al., 2019) and study income differences in life expectancy at age 40. We simulate how our findings would change if we could include younger ages, under two different scenarios for income differences in mortality at these ages. Allocating initial mortality and changes in mortality equally across the rich and poor between the ages of 0-39 has negligible effects. If all early-life mortality is instead allocated to the poor, by construction inequality is initially higher but decreases more since reductions in mortality before age 40 can only be allocated to the poor. The higher initial mortality of the poor at these early ages also implies they have lower survivability. As a result, survivability continues to be a key driver of changes in life expectancy inequality.

Our classification of lifestyle causes of death includes smoking, alcohol, and obesity related deaths not classified as cardiovascular or cancer. If we instead use the causes of death studied in (Case and Deaton, 2015), the conclusions remain largely unchanged.

8 Conclusion

Trends in age-specific mortality rates are informative about the effects of health innovations and health policy on the underlying health status of the rich and poor, while trends in their life expectancy are a relevant measure of the associated welfare effects and the implications for social security policy. Our decomposition and empirical results make clear how focusing on changes in either mortality or life expectancy in isolation can lead to misleading conclusions, and how this can be reconciled by recognizing the way the two measures are linked together through survivability.

References

Arriaga, E. E. (1984). Measuring and explaining the change in life expectancies. *Demography 21*(1), 83–96.

Attanasio, O. P. and T. H. Nielsen (2020). Economic Resources, Mortality and Inequality. *CEBI Working Paper 06/20, University of Copenhagen*.

Auerbach, A. J., K. K. Charles, C. C. Coile, W. Gale, D. Goldman, R. Lee, C. M. Lucas, P. R. Orszag, L. M. Sheiner, B. Tysinger, D. N. Weil, J. Wolfers, and R. Wong (2017). How the Growing Gap in Life Expectancy May Affect Retirement Benefits and Reforms. *Geneva Papers on Risk and Insurance: Issues and Practice 42*(3), 475–499.

Baker, M., J. Currie, and H. Schwandt (2019). Mortality inequality in Canada and the United States: Divergent or convergent trends? *Journal of Labor Economics 37*(S2), S325–S353.

Blinder, A. S. (1973). Wage Discrimination: Reduced Form and Structural Estimates. *The Journal of Human Resources 8*(4), 436–455.

Bor, J., G. H. Cohen, and S. Galea (2017). Population health in an era of rising income inequality: USA, 1980–2015. *Lancet 389*(10077), 1475–1490.

Canudas Romo, V. (2003). Decomposition Methods in Demography. *Max Planck Institute for Demographic Research* (June), 1–178.

Case, A. and A. Deaton (2015). Rising morbidity and mortality in midlife among white non-Hispanic Americans in the 21st century. *Proceedings of the National Academy of Sciences of the United States of America 112*(49), 15078–15083.

Case, A. and A. Deaton (2017). Mortality and morbidity in the 21st century. *BPEA Conference Drafts*, 1–60.

Chetty, R., M. Stepner, S. Abraham, S. Lin, B. Scuderi, N. Turner, A. Bergeron, and D. Cutler (2016). The association between income and life expectancy in the United States, 2001-2014. *JAMA - Journal of the American Medical Association 315*(16), 1750–1766.

Currie, J. and H. Schwandt (2016). Inequality in mortality decreased among the young while increasing for older adults, 1990-2010. *Science 352*(6286), 708–712.

Cutler, D., A. Deaton, and A. Lleras-Muney (2006). The determinants of mortality. *Journal of Economic Perspectives 20*(3), 97–120.

Cutler, D. and S. Kadiyala (2003). Chapter 4: The Return to Biomedical Research: Treatment and Behavioral Effects. In *Measuring the gains from medical research: An economic approach*. Chicago and London: University of Chicago Press.

Cutler, D. M., F. Lange, E. Meara, S. Richards-Shubik, and C. J. Ruhm (2011). Rising educational gradients in mortality: The role of behavioral risk factors. *Journal of Health Economics 30*(6), 1174–1187.

Deaton, A. (2013). *The Great Escape - Health, Wealth, and the Origins of Inequality*. Princeton. Princeton University Press.

Fortin, N., T. Lemieux, and S. Firpo (2011). Decomposition Methods in Economics. In O. Ashenfelter and D. Card (Eds.), *Handbook of Labor Economics*, Volume 4, PART A, Chapter 1, pp. 1–102. Elsevier.

Fuchs, V. R. (1974). *Who Shall Live? Health Economics and Social Choice*. New York, Basic Books.

Goldin, C. and A. Lleras-Muney (2019). XX>XY?: The changing female advantage in life expectancy. *Journal of Health Economics 67*.

Guillot, M. (2011). Period Versus Cohort Life Expectancy. In R. G. Rogers and E. M. Crimmins (Eds.), *International Handbook of Adult Mortality*, Volume 2, Chapter 25, pp. 533–549. Springer.

Haan, P., A. Hammerschmid, and J. Schmieder (2019). Mortality in midlife for subgroups in Germany. *Journal of the Economics of Ageing 14*.

Hederos, K., M. Jäntti, L. Lindahl, and J. Torssander (2017). Trends in Life Expectancy by Income and the Role of Specific Causes of Death. *Economica 85*, 606–625.

Ho, J. Y. and A. S. Hendi (2018). Recent trends in life expectancy across high income countries: Retrospective observational study. *BMJ (Online) 362*.

Honoré, B. E. and A. Lleras-Muney (2006). Bounds in competing risks models and the war on cancer. *Econometrica 74*(6), 1675–1698.

Jayachandran, S., A. Lleras-Muney, and K. V. Smith (2010). Modern medicine and the twentieth century decline in mortality: Evidence on the impact of sulfa drugs. *American Economic Journal: Applied Economics 2*(2), 118–146.

Jdanov, D. A., V. M. Shkolnikov, A. A. van Raalte, and E. M. Andreev (2017). Decomposing Current Mortality Differences Into Initial Differences and Differences in Trends: The Contour Decomposition Method. *Demography 54*(4), 1579–1602.

Kinge, J. M., J. H. Modalsli, S. Øverland, H. K. Gjessing, M. C. Tollånes, A. K. Knudsen, V. Skirbekk, B. H. Strand, S. E. Håberg, and S. E. Vollset (2019). Association of Household Income with Life Expectancy and Cause-Specific Mortality in Norway, 2005-2015. *JAMA - Journal of the American Medical Association 321*(19), 1916–1925.

Kitagawa, E. M. (1955). Components of a Difference Between Two Rates. *Journal of the American Statistical Association 50*(272), 1168–1194.

Kreiner, C. T., T. H. Nielsen, and B. L. Serena (2018). Role of income mobility for the measurement of inequality in life expectancy. *Proceedings of the National Academy of Sciences of the*

United States of America 115(46), 11754–11759.

Likosky, D. S., J. Van Parys, W. Zhou, W. B. Borden, M. C. Weinstein, and J. S. Skinner (2018). Association between medicare expenditure growth and mortality rates in patients with acute myocardial infarction: A comparison from 1999 through 2014. *JAMA Cardiology 3*(2), 114–122.

Lleras-Muney, A. (2005). The relationship between education and adult mortality in the United States. *Review of Economic Studies 72*(1), 189–221.

Mackenbach, J. P., J. R. Valverde, B. Artnik, M. Bopp, H. Brønnum-Hansen, P. Deboosere, R. Kalediene, K. Kovács, M. Leinsalu, P. Martikainen, G. Menvielle, E. Regidor, J. Rychtaríkova, M. Rodriguez-Sanz, P. Vineis, C. White, B. Wojtyniak, Y. Hu, and W. J. Nusselder (2018). Trends in health inequalities in 27 European countries. *Proceedings of the National Academy of Sciences of the United States of America 115*(25), 6440–6445.

Marmot, M., J. Allen, P. Goldblatt, T. Boyce, D. McNeish, M. Grady, and I. Geddes (2010). Fair society, healthy lives (The Marmot Review): Strategic Review of Health Inequalities in English post-2010. Technical report.

Montez, J. K., A. Zajacova, M. D. Hayward, S. H. Woolf, D. Chapman, and J. Beckfield (2019). Educational Disparities in Adult Mortality Across U.S. States: How Do They Differ, and Have They Changed Since the Mid-1980s? *Demography 56*(2), 621–644.

Moscelli, G., L. Siciliani, N. Gutacker, and R. Cookson (2018). Socioeconomic inequality of access to healthcare: Does choice explain the gradient? *Journal of Health Economics 57*, 290–314.

Oaxaca, R. (1973). Male-Female Wage Differentials in Urban Labor Markets. *International Economic Review 14*(3), 693–709.

Poterba, J. M. (2014). Retirement security in an aging population. *American Economic Review 104*(5), 1–30.

Ruhm, C. (2018). Deaths of Despair or Drug Problems? *NBER Working Paper No. 24188*.

Snyder, S. E. and W. N. Evans (2006). The Effect of Income on Mortality : Evidence from the Social Security Notch. *Review of Economics and Statistics 88*(3), 482–495.

van den Berg, G. J., U. G. Gerdtham, S. von Hinke, M. Lindeboom, J. Lissdaniels, J. Sundquist, and K. Sundquist (2017). Mortality and the business cycle: Evidence from individual and aggregated data. *Journal of Health Economics 56*, 61–70.

Waldron, H. (2007). Trends in mortality differentials and life expectancy for male social security-covered workers, by socioeconomic status. *Social Security Bulletin 67*(3), 1–28.

Tables

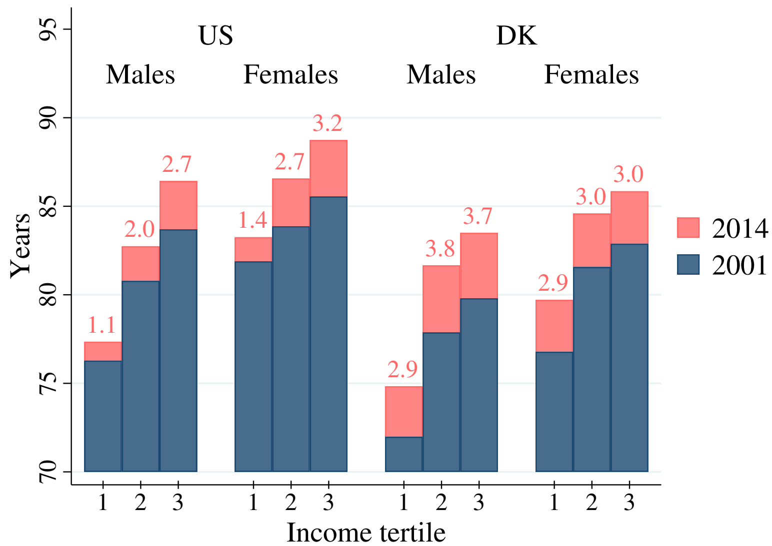
Table 1: Effect of mortality change at age 60 on life expectancy of 40-year-old males

	ΔMortality ΔM_{60} (1)	Survival S_{60} (2)	Remain. LE R_{60} (3)	Survivability X_{60} (4)=(2)x(3)	ΔLife Exp ΔLE_{40} (5)=(1)x(4)
United States					
Poor	-0.002	0.84	19.0	16.0	0.032
Rich	-0.002	0.95	23.5	22.3	0.045
Denmark					
Poor	-0.003	0.78	15.6	12.2	0.037
Rich	-0.003	0.94	19.6	18.4	0.055

Notes: The table shows the effect of the observed same change in the mortality rate of rich and poor 60 year old males on life expectancy at age 40. The table reports: (1) changes in mortality rates during 2001-2014, (2) survival probabilities, S_{60}, up to age 60, (3) remaining life expectancy after age 60, R_{60}, (4) survivability at age 60, $X_{60} \equiv S_{60} \cdot R_{60}$, and (5) the effect on life expectancy at age 40 from the changes in mortality rates at age 60. The change in life expectancy in column (5) is computed from the first-order approximation in equation (2). Survival probabilities and remaining life expectancies are calculated using the definitions reported below equation (2) and measured in 2001 in accordance with the first-order approximation. To match the numbers in Figure 2, which plots mortality rate changes and survivability by 5-year age bins, all numbers are calculated as the average of individuals aged 60-64 years.

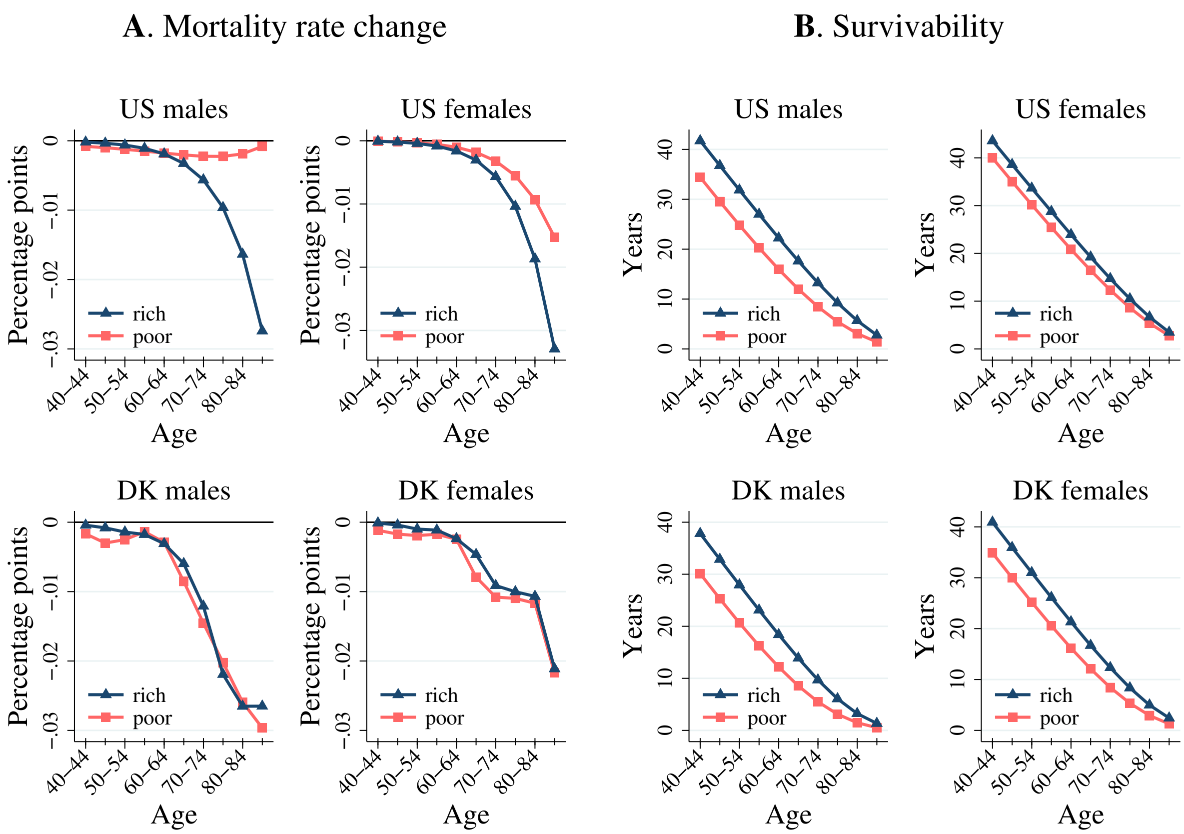
Figures

Figure 1: Life expectancy at age 40 by income class, US and Denmark



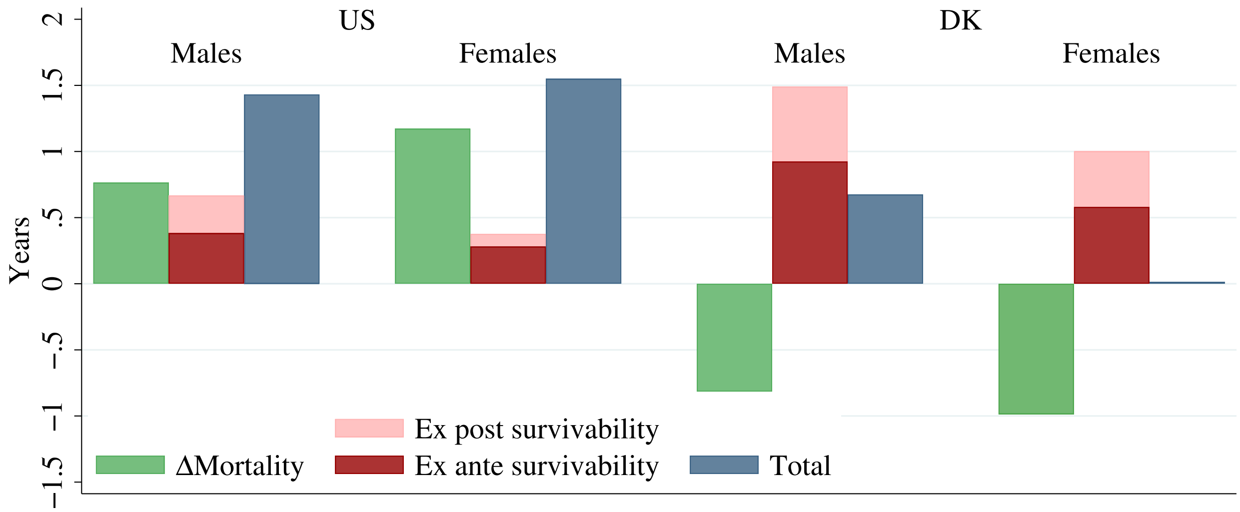
Notes: The figure shows the average life expectancy at age 40 in the US and Denmark by income tertiles in 2001 and 2014 for males and females. The numbers above the bars denote the increase from 2001 to 2014. The figure is constructed by computing period life expectancy for each tertile group using equation (1) and data on mortality rates by age and income class. The higher life expectancy for the US compared to Denmark reflects that the US data excludes individuals with zero earnings, while the Danish data covers the full population. Conclusions are unchanged if we impose a similar restriction on the Danish data to make estimates comparable (see Appendix A.2). The data is described in Section 3 and Appendix A.1.

Figure 2: Mortality changes and survivability, by age and income class



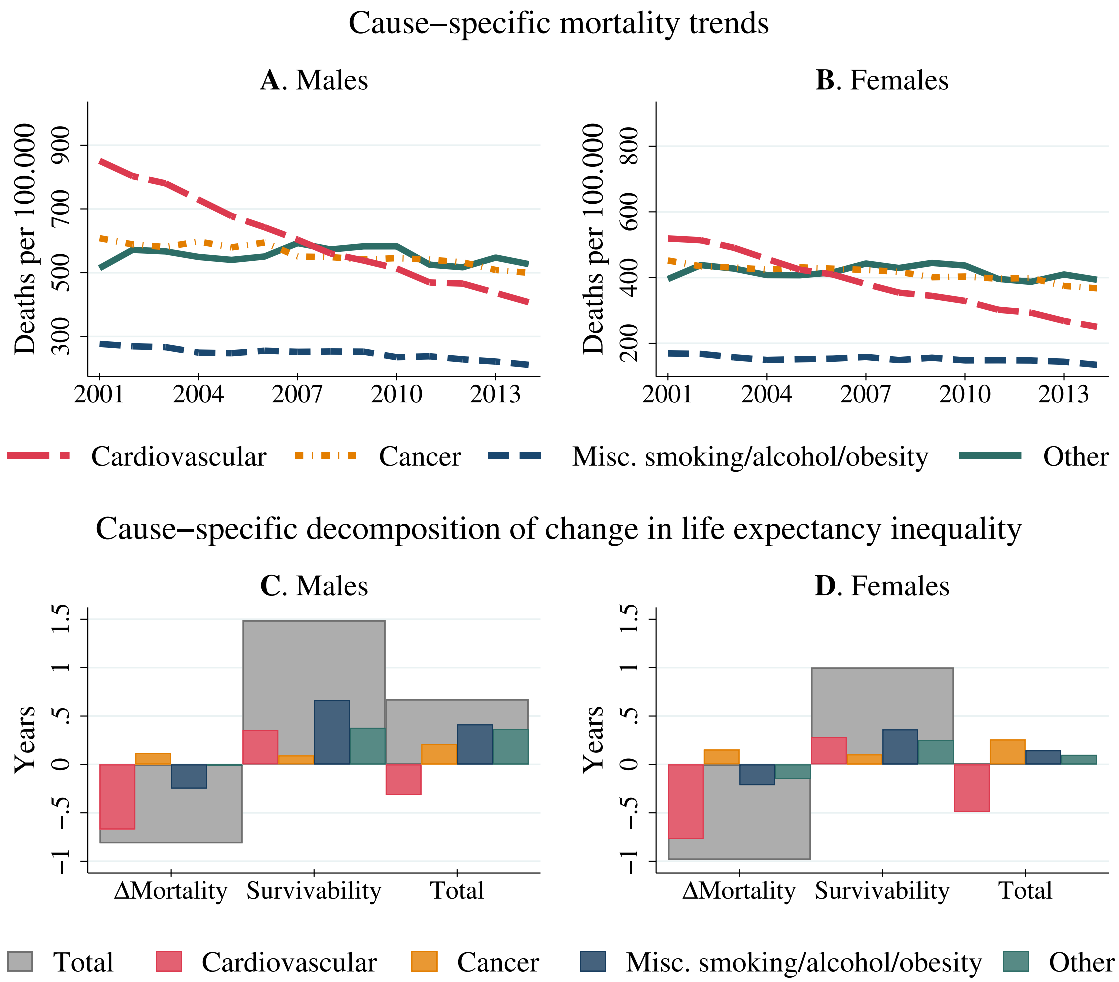
Notes: Panel A plots changes in mortality rates from 2001 to 2014 by 5-year age bins for poor (tertile 1) and rich (tertile 3) males and females in the US and Denmark. The figures are constructed by computing average changes in mortality rates within 5-year age groups. Panel B plots survivability by 5-year age bins for the same groups. These figures are constructed by first using the definition in equation (2) and data on mortality rates to compute survivability, X_a, in 2001 and then averaging across 5-year age groups.

Figure 3: Change in life expectancy gap between rich and poor at age 40 decomposed into differential mortality rate changes and differences in survivability.



Notes: To create the figure, we use the decomposition formulas in equations (3) and (4) and the data on mortality rates from the US and Denmark to compute the change in life expectancy inequality between the rich and poor during 2001-2014 and to decompose this into (1) contributions from differential changes in mortality rates and (2) differences in (ex ante and ex post) survivability.

Figure 4: Cause-specific decomposition of change in life expectancy inequality, DK



Notes: Panels A and B plot age-standardized mortality per 100,000 individuals in Denmark by cause of death for the years 2001 to 2014 for males and females. The age standardization uses the US standard population from the World Health Organization, downloaded from `https://seer.cancer.gov/stdpopulations/stdpop.singleages.html`. Panels C and D show the cause-specific decomposition of changes in inequality in life expectancy for males and females measured at age 40 and computed using equation (5) based on cause and income-specific mortality rates in the Danish data. Miscellaneous smoking/alcohol/obesity-related diseases consist of COPD (ICD-10: J40-J44), alcohol abuse (F10), alcoholic liver disease and cirrhosis (K70, K74), alcohol-induced chronic pancreatitis (K86.0), diabetes (E10-E14), obesity (E66).

Online Appendix for
Understanding the Rise in Life Expectancy Inequality

Gordon B. Dahl, Claus Thustrup Kreiner, Torben Heien Nielsen, Benjamin Ly Serena

Contents

A.1 Data

US. We use publicly available data from Chetty et al. (2016) on income-specific mortality in the US during the years 2001-2014. The income-specific mortality rates are estimated by Chetty et al. (2016) using administrative tax data covering the entire US population. To minimize reverse causality, mortality rates are estimated based on income measured two years prior. The income concept is defined as the sum of household gross income minus Social Security and disability benefits. For individuals who do not file taxes, the income concept includes wage earnings and unemployment benefits. The authors exclude observations with zero income. This restriction drops about 9 percent of the population. As a large fraction of individuals with zero income are disability insurance recipients and these have above-normal rates of mortality, the computed life expectancies are overstated compared to the population averages in the US. After the public retirement age earnings become a poor measure of socioeconomic status. The authors therefore impute mortality rates using Gompertz approximations, which assumes mortality is log-linear in age, $log(M_a) = \alpha + \beta * Age$ (see Chetty et al., 2016). After age 90, at which point the Gompertz approximation is less accurate, the authors replace age and income-specific mortality rates with age-specific, population-wide mortality rates. In addition, Chetty et al. (2016) adjust mortality rates for differences in the composition of ethnicity across income groups.

Denmark. We use Danish administrative data provided by Statistics Denmark and covering the period 2001-2014 for the universe of Danish residents. We link the population registers, BEF and FAIN, containing information on sex and age, with the income register, IND, and the death registers, DODAARS and DODAASG, using personal identifiers (CPR number). Following previous work (Kreiner et al., 2018), we measure income as household income net of universal transfers: Income = PERINDKIALT − QPENSNY − KONTHJ − ANDOVERFORSEL, where PERINDKIALT is total income, QPENSNY is public pensions (disability pension, public retirement pension), KONTHJ is cash assistance, and ANDOVERFORSEL is other universal public transfers. Household income is defined as the mean income of the individual and the spouse (if the individual of interest is either married or cohabiting). Following the previous literature (Chetty et al., 2016; Kreiner et al., 2018), we measure income two years before we measure mortality to reduce the importance of reverse causality. Because the data covers the entire population of Denmark, the estimated life expectancies line up closely with official estimates from Statistics Denmark. The income measure we use includes payouts from private and employer-based pension accounts, which are tightly linked to previous labor market earnings. For this reason, our income concept remains a good measure of socioeconomic status after retirement. See Kreiner et al. (2018) for a validation of the income measure and the stability of income ranks around retirement.

Below, in the section titled 'Comparability of results for US and Denmark', we discuss how differences in data sources affect comparisons between Denmark and the US and present estimates for Denmark based on data restrictions similar to those in the US data.

Construction of dataset for the main analysis. The main results are based on age-specific mortality rates by income tertiles in 2001 and 2014. This is constructed by first computing annual mortality rates by age, sex, year, and income tertile (within cohort, sex, and year) for the ages 40-100 and years 2001-2014. To maximize precision, we then use all years during 2001-2014 to run regressions of age-specific mortality rates on linear trends and use the predicted age-specific mortality rates in 2001 and 2014 rather than the actual mortality rates for these two years. For Denmark, we have information on causes of death. Therefore, we also compute four different cause-specific mortality rates; cardiovascular (ICD-10 codes: I00-I78), smoking/alcohol/obesity-related (ICD-10 codes: COPD: J40-J44, alcohol abuse: F10, alcoholic liver disease and cirrhosis: K70, K74, alcohol-induced chronic pancreatitis K86.0, diabetes: E10-E14, obesity: E66), cancer (ICD-codes: C00-C97, D00-D09), and other (all remaining ICD-10 codes). We follow the same procedure as for all-cause mortality rates and predict cause-specific mortality rates in 2001 and 2014 using regressions with linear trends through all years during 2001-2014.

A.2 Comparability of results for the US and Denmark

The US income data differ in important ways from the Danish data. While the Danish data covers the entire population, the US data does not include individuals with zero income, many of whom are disability insurance recipients. This sample restriction drops 9% of the US population and 32% of deaths. Therefore, mortality rates for the US are underestimated and life expectancies are overestimated compared to population averages. In Figure A.1, we exclude individuals with zero or negative income and disability insurance recipients from the Danish data, to make results for Denmark and the US more comparable. These restrictions drop 9% of the Danish population and 36% of deaths. After the public retirement age, disability insurance recipients receive a public pension and cannot be identified in the data. Instead, we calculate the share of individuals in each income group dropped by our restrictions before retirement (age 60) and drop the same share in post retirement ages.

Figure A.1A plots the average life expectancy in Denmark and the US by income tertiles in 2001 and 2014 for males and females. As most of the disability insurance recipients have high mortality and low income, implementing the US sample restrictions increases average life expectancy and reduces inequality in Denmark. For example, the difference between males in the bottom and top tertile of income in 2001 is 5 years rather than 8 years. As shown in Figure A.1B, the rise in life expectancy inequality among Danish males from 2001 to 2014 is also smaller; 0.5 years instead of 0.7 years. However, the sample restriction does not change our main conclusions. Declines in mortality over time are larger for the poor than the rich, and the rise in life expectancy inequality in Denmark is driven entirely by differences in survivability.

A.3 Derivation of decomposition formulas

Equation (2) in the main text is derived by differentiation of the definition for life expectancy in equation (1) in the main text. By differentiating this expression with respect to the mortality rate M_a at age a, we get the first-order approximation:

$$\Delta LE_{\underline{a}} \approx -\Delta M_a \prod_{i=\underline{a}}^{a-1}(1-M_i)\sum_{j=a}^{\infty}\prod_{i=a+1}^{j}(1-M_i) \approx -\Delta M_a \prod_{i=\underline{a}}^{a-1}(1-M_i)\left(1+\sum_{j=a+1}^{\infty}\prod_{i=a+1}^{j}(1-M_i)\right)$$

where we have used the definition $\prod_{i=a+1}^{a}(1-M_i) \equiv 1$. Next, we define

$$S_a \equiv \prod_{i=\underline{a}}^{a-1}(1-M_i), \quad R_a \equiv 1+\sum_{j=a+1}^{\infty}\prod_{i=a+1}^{j}(1-M_i) \tag{A.1}$$

where S_a is the probability of survival from age $\underline{a}$ to age a and R_a is remaining life expectancy if surviving age a, in which case the individual lives one extra year for sure and from age $a+1$ can expect to live the additional years given by the last term in the definition. This implies that the effect of a change in the mortality rate M_a at age a may be written as:

$$\Delta LE_{\underline{a}} \approx -\Delta M_a \cdot X_a \text{ where } X_a \equiv S_a \cdot R_a. \tag{A.2}$$

The effect of a change in all mortality rates $\{M_i\}_{\underline{a}}^{\infty}$ can be found by total differentiation of definition (1) in the main text, which corresponds to a simple aggregation of equation (A.2). This gives

$$\Delta LE_{\underline{a}} \approx -\sum_{a=\underline{a}}^{\infty}\Delta M_a \cdot X_a. \tag{A.3}$$

This is equation (2) in the main text. We proceed by deriving changes in life expectancy inequality over time. We study life expectancy inequality between a poor group (LE^p) and a rich group (LE^r) with equally many individuals in each group (bottom one-third versus top one-third). Using equation (A.3), the change in life expectancy inequality over time equals:

$$\Delta LE_{\underline{a}}^r - \Delta LE_{\underline{a}}^p \approx -\sum_{a=\underline{a}}^{\infty}\left(\Delta M_a^r \cdot X_a^r - \Delta M_a^p \cdot X_a^p\right). \tag{A.4}$$

Equation (A.3) may be rewritten as

$$\Delta LE_{\underline{a}}^k \approx -\sum_{a=\underline{a}}^{\infty}\Delta M_a^k \cdot \overline{X_a} - \sum_{a=\underline{a}}^{\infty}\overline{\Delta M}_a \cdot \left(X_a^k - \overline{X_a}\right) - \sum_{a=\underline{a}}^{\infty}\left(\Delta M_a^k - \overline{\Delta M}_a\right)\cdot\left(X_a^k - \overline{X_a}\right), \quad k=r,p$$

where superscript k denotes the income class. We insert this expression in equation (A.4) along with the definition of the averages $\overline{X_a} = \frac{1}{2}\left(X_a^r + X_a^p\right)$ and $\overline{\Delta M}_a = \frac{1}{2}\left(\Delta M_a^r + \Delta M_a^p\right)$:

$$\begin{aligned}\Delta LE_{\underline{a}}^r - \Delta LE_{\underline{a}}^p \approx & -\sum_{a=\underline{a}}^{\infty}\left[\left(\Delta M_a^r - \Delta M_a^p\right)\cdot\overline{X_a} + \overline{\Delta M}_a \cdot \left(X_a^r - X_a^p\right)\right] \\ & -\sum_{a=\underline{a}}^{\infty}\frac{1}{4}\left(\Delta M_a^r - \Delta M_a^p\right)\cdot\left(X_a^r - X_a^p\right) + \sum_{a=\underline{a}}^{\infty}\frac{1}{4}\left(\Delta M_a^p - \Delta M_a^r\right)\cdot\left(X_a^p - X_a^r\right)\end{aligned}$$

The last two (covariance) terms cancel out because the two groups are of the same size. Thus, we arrive at equation (3) in the main text.

We derive equation (4) in the main text by splitting equation (3) into contributions from ex ante and ex post survivability. We use the definition of survivability, $X_a \equiv S_a \cdot R_a$, and follow the same steps as in the derivation of equation (3) described above.

$$\begin{aligned}\Delta LE^r_{\underline{a}} - \Delta LE^p_{\underline{a}} &\approx -\sum_{a=\underline{a}}^{\infty} \Big[(\Delta M^r_a - \Delta M^p_a) \cdot \overline{X_a} + \overline{\Delta M}_a \cdot (S^r_a \cdot R^r_a - S^p_a \cdot R^p_a)\Big] \\ &\approx -\sum_{a=\underline{a}}^{\infty} \Big[(\Delta M^r_a - \Delta M^p_a) \cdot \overline{X_a} + \overline{\Delta M}_a \cdot \Big(\overline{S_a} \cdot (R^r_a - R^p_a) + \overline{R_a} \cdot (S^r_a - S^p_a)\Big)\Big]\end{aligned}$$

Empirically, we set mortality rates to 1 after age 100, as there are too few survivors left in the subsequent ages to estimate income-specific mortality.

A.4 Accuracy of first order approximation

The age decomposition formula (2) in the main text is a first-order approximation of the change in life expectancy over time. Figure A.2A shows the difference between the first order approximation and the actual changes in life expectancy for the rich and poor. Overall, the first order approximation provides increases which are too small over time in life expectancy and life expectancy inequality. However, the errors are relatively small. For US females, which has the largest approximation error, the actual increase in life expectancy inequality during 2001-2014 is 1.8 years, while the first order approximation gives 1.5 years.

Often researchers use an Arriaga age decomposition (Arriaga, 1984). In addition to the first-order approximation (2) in the main text, this approximation includes an extra term, an interaction effect, equal to $\Delta R_a \cdot \Delta M_a \cdot S_a$. The interaction effect cannot be attributed to any particular age, but reduces the approximation error of the age decomposition. Figure A.2B shows that our decomposition of trends in life expectancy inequality into differential changes in mortality rates of rich and poor and differences in survivability is more or less unchanged if we include the interaction effect in the age decomposition. Differences in survivability still explain all of the increase in life expectancy inequality in Denmark and between 25-50% of the increase in the US.

A.5 Derivation of cause-specific decomposition

This section describes how we estimate the cause-specific contributions to the change in life expectancy inequality over time and its decomposition into mortality rate changes and survivability in Section 5 of the paper. We start from equation (4) in the main text, which decomposes changes in life expectancy inequality into contributions from changes in mortality rates, ex ante survivability, and ex post survivability. We decompose each of the three terms separately by causes of death denoted by the index c.

Mortality rate changes by cause of death. As cause-specific mortality rates sum to total mortality, it is straightforward to decompose mortality rate changes by cause c:

$$\sum_{a=\underline{a}}^{\infty} \overline{X}_a \cdot (\Delta M_a^p - \Delta M_a^r) = \sum_{a=\underline{a}}^{\infty} \overline{X}_a \cdot \sum_c \left(\Delta M_{a,c}^p - \Delta M_{a,c}^r\right) \tag{A.5}$$

Ex ante survivability by cause of death. To decompose ex ante survivability by cause of death, we rewrite income differences in the probability of surviving to a certain age a, i.e., $S_a^p - S_a^r$, as the sum of differences in cause-specific survival probabilities $S_a^p - S_a^r = \sum_c (S_{a,c}^p - S_{a,c}^r)$, where the cause-specific survival probabilities reflect the probability that a person does not die from that particular disease between age $\underline{a}$ and a. We insert this into the expression for ex ante survivability :

$$\sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{R_a} \cdot (S_a^p - S_a^r) = \sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{R_a} \cdot \sum_c (S_{a,c}^p - S_{a,c}^r) \tag{A.6}$$

Ex post survivability by cause of death. Decomposing ex post survivability by cause of death is analogous to decomposing differences in life expectancy by cause, which is standard. We rewrite differences in remaining life expectancy, R_a, as the sum of differences in survival probabilities in succeeding ages:

$$R_a^p - R_a^r = \sum_{j=a+1}^{\infty} \left(\prod_{i=a+1}^{j} (1 - M_i^p) - \prod_{i=a+1}^{j} (1 - M_i^r) \right) = \sum_{j=a+1}^{\infty} \left(\tilde{S}_j^p - \tilde{S}_j^r\right) \tag{A.7}$$

where $\tilde{S}_j \equiv \prod_{i=a+1}^{j} (1 - M_i)$ is the probability of surviving from age $a + 1$ to past age j. As in equation (A.6), we can rewrite differences in survival as the sum of differences in cause-specific survival probabilities $\tilde{S}_j^p - \tilde{S}_j^r = \sum_c (\tilde{S}_{j,c}^p - \tilde{S}_{j,c}^r)$. Furthermore, we define $\tilde{Q}_{a,c}^k \equiv \sum_{j=a+1}^{\infty} (1 - \tilde{S}_{j,c}^k)$ for $k = r, p$, which denotes expected lost life measured in years from age a due to the risks of dying of cause c, where $\tilde{Q}_a = 100 - a - R_a$ and 100 is the highest attainable life expectancy as, empirically, we set mortality rates to 1 after this age. Inserting $\tilde{Q}_a$ into the expression for ex post survivability yields:

$$\sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{S_a} \cdot (R_a^p - R_a^r) = \sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot \overline{S_a} \cdot \sum_c (\tilde{Q}_{j,c}^r - \tilde{Q}_{j,c}^p) \tag{A.8}$$

where the cause-specific contribution to differences in remaining life expectancy, $(R_{j,c}^p - R_{j,c}^r)$ is given by the cause-specific contribution to lost life years.

A.6 Decomposition by educational attainment

Throughout the main text we focus on income differences in life expectancy. Education is another often-used measure of socioeconomic status. Figure A.3 shows our full cause-specific decomposition results using education instead of income for both the US and Denmark. For the US, we replicate estimates of mortality by education in the US by Case and Deaton (2017) using data from the National Vital Statistics and the March Current Population Survey. As in Case and Deaton (2017), we consider the difference in life expectancy between individuals with high school or less, some college, and BA or more. In Denmark, where data on highest completed education is available from administrative records, we follow Mackenbach et al. (2018) and define low, middle, and high education groups using International Standard Classification of Education 1997 (ISCED) codes; low: 0-2, middle: 3-4, high: 5+. To account for changes in the composition of education groups over time, we hold the share of the population in each education group fixed. Hence, as the share of the population in low education groups decreases over time, we randomly allocate some individuals in the middle education group to the low education group. This is similar to using ranks, as we do for income.

In both Denmark and the US, education information is only available for individuals below age 80. Therefore, we impute mortality from age 80 to 100 using Gompertz approximations. Following Chetty et al. (2016), we replace these approximations with population-wide mortality rates after age 90, because the Gompertz approximation is less accurate after this age.

Both the Danish and US data contain information on cause-specific mortality, allowing us to conduct cause-specific decompositions. However, because we impute mortality rates by education after age 80, information on cause of death by education is not available in these ages. To deal with this issue, we assume that the relative importance of different diseases for total mortality across education groups is fixed at the age 80 level. Hence, if the share of deaths from cardiovascular disease within an education group is twice as large for the low-educated as the high-educated at age 80, we assume that this is also the case at age 81 and so forth.[1] After age 90, all education groups are assigned population-wide, cause-specific mortality rates.

Figure A.3A and B show the development in cause-specific mortality in the US from 2001 to

[1] Assume the share of deaths from a given disease is proportional across education groups (low=l, mid=m, high=h):

$$\frac{D^l_{a,c}}{D^l_a} = x \cdot \frac{D^m_{a,c}}{D^m_a}, \quad \frac{D^l_{a,c}}{D^l_a} = y \cdot \frac{D^h_{a,c}}{D^h_a} \tag{A.9}$$

where $D_{a,c}$ is cause-specific mortality and the values of x and y are measured at age 80 and assumed constant thereafter. The sum of education-specific mortality rates must sum to total mortality, $N^l_a \cdot D^l_{a,c} + N^m_a \cdot D^m_{a,c} + N^h_a \cdot D^h_{a,c} = D_{a,c}$ where N^e_a denotes the share of individuals alive at age a that belong to education group e. We can then isolate $D^l_{a,c}$ as: $D^l_{a,c} = D_{a,c}\left(\frac{D^m_a \cdot N^m_a}{D^l_a \cdot x} + N^l_a + \frac{D^h_a \cdot N^h_a}{D^l_a \cdot y}\right)^{-1}$. $D^m_{a,c}$ and $D^h_{a,c}$ follow from equation (A.9). This imputation does not ensure that the sum of the cause-specific mortality rates sum to aggregate mortality within each education group. However, it is very close empirically. We fix the small difference by multiplying cause and education-specific mortality rates with $\frac{D^e_a}{\sum_c D^e_{a,c}}$.

2014 using data from the WHO, and are similar to Figure 4A and B for Denmark in the main text. In line with the numbers for Denmark, the figure for the US shows a large drop in cardiovascular deaths over time and this is the main driver of the increase in life expectancy during 2001-2014.

Figure A.3C shows the decomposition results for the US and Denmark using educational attainment. Overall, the results for education are similar to those for income although the magnitudes differ somewhat. Inequality in life expectancy between the low and high-educated in the US increased by around 1 year during 2001-2014 for both males and females. During this period, changes in mortality rates have favored the highly educated, but this explains a smaller fraction of the increase in inequality in life expectancy across education groups compared to when we used income groups to measure socioeconomic status. Differences in mortality rate changes account for about 10% of the increase in inequality for males and about 60% for females. Consequently, the remaining 90% for males and 40% for females can be attributed to differences in survivability across education groups. The education results for Denmark are also similar to the income results. Changes in mortality rates over time have favored the low-educated, but because of differences in survivability, the improvements in mortality end up having a smaller impact on life expectancy for the low-educated than for the high-educated.

The cause-specific decomposition results for education are also similar to those for income. For both countries and both sexes, the drop in mortality rates related to cardiovascular disease reduced the gap in life expectancy between the low and high-educated. Nevertheless, total inequality in life expectancy increased. A key reason is the difference between the low and high-educated in survivability where differences in lifestyle-related causes of death play a major role.

A.7 Inequality trends in other Western European countries

Our main results show that mortality trends favor the poor in Denmark, but the rich in the US. In Figure A.4, we apply our decomposition to education inequalities in life expectancy across a number of European countries. We estimate these using published mortality rates by education groups from Mackenbach et al. (2018). The education groups are based on the same International Standard Classification of Education 1997 (ISCED) codes we use for Denmark in the previous section of this appendix. The data contains information on one-year mortality rates by 5-year age groups from age 40 to 75 and for various years between 1997 and 2013. The years available in the data differ across countries and therefore we report yearly changes in life expectancy. As the data is only available up until age 75 and only for 5-year age groups, we predict mortality rates using Gompertz approximations. After age 90, we apply population-wide mortality rates to all education groups, obtained from life tables estimated by the WHO (`http://ghdx.healthdata.org/gbd-results-tool`). Because the Gompertz approximations are based on 5-year age groups and only 7 observations per country, these results should be interpreted with caution. Also, when estimating life expectancies by education in the previous section, we account for changes in the composition of education over time. We cannot do this in the present analysis because we do not have information on the education shares over time in these countries.

Despite these caveats, the decomposition results across the nine Western European countries in Figure A.4 largely mimic the results for Denmark. Among males, all countries but one (Spain) have seen larger mortality rate reductions among the low-educated than among the high-educated. In six of the remaining eight countries, the favorable mortality trends for the low-educated are associated with increasing inequality in life expectancy due to differences in survivability. The results for females are more mixed. In seven out of nine countries, mortality trends have favored the low-educated, but in most of these cases this has led to reductions in life expectancy inequality. In two Nordic countries, Finland and Sweden, mortality rate changes have favored the high-educated and so have changes in life expectancy.

A.8 Relative changes in mortality rates and life expectancy

Our main analysis focuses on absolute changes in mortality rates as is commonly done in the literature (Currie and Schwandt, 2016; Case and Deaton, 2015; Kinge et al., 2019). One may ask whether differential relative changes in mortality rates between rich and poor map one-to-one into differential changes in their life expectancy. Below we show that this is not the case. Differences in survivability are still important and life expectancy may rise more for the rich than the poor even when relative changes in mortality rates are the same for both groups. We first augment our decomposition formula to study differences in relative changes in mortality rates. Using $\Delta M_a \equiv \frac{\Delta M_a}{M_a} \cdot M_a$, we may rewrite equation (3) in the main text as

$$\Delta LE_{\underline{a}}^r - \Delta LE_{\underline{a}}^p = \sum_{a=\underline{a}}^{\infty} \overline{X_a} \cdot \left(\frac{\Delta M_a^p}{M_a^p} \cdot M_a^p - \frac{\Delta M_a^r}{M_a^r} \cdot M_a^r \right) + \sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot (X_a^p - X_a^r)$$

By means of the method described in Section A.3 above, this can be further decomposed into changes in relative mortality rates and differences in initial mortality rates:

$$\Delta LE_{\underline{a}}^r - \Delta LE_{\underline{a}}^p = \sum_{a=\underline{a}}^{\infty} \overline{X_a} \cdot \overline{M_a} \cdot \left(\frac{\Delta M_a^p}{M_a^p} - \frac{\Delta M_a^r}{M_a^r} \right) + \sum_{a=\underline{a}}^{\infty} \overline{X_a} \cdot \overline{\left(\frac{\Delta M_a}{M_a} \right)} \cdot (M_a^p - M_a^r) + \sum_{a=\underline{a}}^{\infty} \overline{\Delta M_a} \cdot (X_a^p - X_a^r) \tag{A.10}$$

The first term in the decomposition can be interpreted as the change in life expectancy inequality that would occur if relative changes in mortality rates differed across income groups, but initial mortality rates were the same. The second term reflects that, if the number of individuals surviving to a given age is the same across income groups, the same percentage drop in mortality will increase the number of survivors at that age more for the poor than the rich, because the poor have higher baseline mortality. Hence, if baseline mortality is twice as large for the poor, the same percentage drop in mortality will lead to twice as many new survivors at that age. The third term, which is just the standard survivability term, reflects that fewer among the poor will survive to benefit from a reduction in mortality at a given age (ex ante survivability) and that for each new survivor at that age, the remaining life expectancy is lower for the poor than the rich (ex post survivability).

Importantly, this decomposition shows that regardless of whether we study absolute or relative mortality rate changes, we cannot tell whether inequality in life expectancy is increasing or decreasing on the basis of mortality rate changes alone. Even equal relative changes in mortality rates across income groups can lead to increasing life expectancy inequality because of differences in survivability. Table A.1 provides an empirical example. For males in Denmark, column (1) shows that mortality at age 85 has dropped by approximately 18% for both the rich and the poor. With initial mortality rates of 0.17 for the poor and 0.14 for the rich, see column (2), this means that among those surviving to age 85, the share of individuals surviving beyond age 85 has increased more for the poor (3 percentage points) than the rich (2.5 percentage points). In isolation, this reduces inequality. However, as shown in column (3) of Table A.1, only 12% of the poor survive long enough to benefit from the reduction in mortality at age 85 versus 28% of the rich. In addition, for those who now survive beyond age 85, column (4) shows that the remaining life expectancy is 4.3 years for the poor compared to 4.8 years for the rich. These differences in survivability, reported in column (5), imply that the same relative change in mortality rates across income groups leads to twice the increase in life expectancy for the rich compared to the poor, as shown in column (6). Column (7) shows that the percentage increase in life expectancy is also highest for the rich.

A.9 Disability-adjusted life expectancy

Life expectancy is a mortality-based summary measure of health in a population. This does not account for the quality of life or burden of disease. For this reason, researchers and health organizations often compute health expectancies or disability-free life expectancies, which put lower weight on life years lived with disability. Figure A.6 displays our main decomposition results using disability-free life expectancy $DFLE$. We do this by applying a weight (W_a), which measures the value of a year of life at each age. This is equivalent to the often-used Sullivan method (Sullivan, 1971).

$$
\begin{aligned}
DFLE_{\underline{a}} =& \underline{a} + W_{\underline{a}} \cdot (1 - M_{\underline{a}}) + W_{\underline{a}+1} \cdot (1 - M_{\underline{a}}) \cdot (1 - M_{\underline{a}+1}) + ... \\
=& \underline{a} + \sum_{j=\underline{a}}^{\infty} W_j \prod_{i=\underline{a}}^{j} (1 - M_i)
\end{aligned} \tag{A.11}
$$

Differentiating with respect to mortality at a given age a yields:

$$
\begin{aligned}
\Delta DFLE_{\underline{a}} =& - \Delta M_a \cdot \prod_{i=\underline{a}}^{a-1} (1 - M_i) \cdot \sum_{j=a}^{\infty} W_j \prod_{i=a+1}^{j} (1 - M_i) \\
=& - \Delta M_a \cdot \prod_{i=\underline{a}}^{a-1} (1 - M_i) \cdot \left(W_a + \sum_{j=a+1}^{\infty} W_j \prod_{i=a+1}^{j} (1 - M_i) \right) \\
=& - \Delta M_a \cdot \hat{X}_a
\end{aligned} \tag{A.12}
$$

where $\hat{X}_a \equiv S_a \cdot \hat{R}_a$ is disability-free survivability and $\hat{R}_a \equiv W_a + \sum_{j=a+1}^{\infty} W_j \prod_{i=a+1}^{j} (1 - M_i)$ is disability-free remaining life expectancy at age a. From (A.12), we follow the steps described in Section A.3 and decompose changes in inequality in disability-free life expectancy into contributions from differences in mortality rate changes and survivability:

$$
\Delta DFLE_{\underline{a}}^{r} - \Delta DFLE_{\underline{a}}^{p} = \sum_{a=\underline{a}}^{\infty} \left[(\Delta M_a^p - \Delta M_a^r) \cdot \overline{\hat{X}_a} + \overline{\Delta M_a} \cdot \left(\hat{X}_a^p - \hat{X}_a^r \right) \right] \tag{A.13}
$$

To implement this decomposition, we need estimates of the disease weights W_a. We base these on WHO estimates of years lived with disability:

$$
W_a = 1 - \frac{YLD_a}{P_a - D_a} \tag{A.14}
$$

where YLD_a is years lived with disability (summing over the entire population), P_a is the population and D_a is the number of deaths at age a. We obtain these three components by 5-year age groups from WHO's Global Burden of Disease database, `http://ghdx.healthdata.org/gbd-results-tool`. The weights measure the share of years lived at age a, $P_a - D_a$, that are spent without disability. WHO's years lived with disability estimates are based on the prevalence of certain diseases and the associated loss in life satisfaction, see `https://www.who.int/healthinfo/global_burden_disease/metrics_daly/en/`. Note that the weights we calculate are population wide since we do not have information on burden of disease by income.

Figure A.6A plots the weights by age for Danish and US males and females in 2001. In both countries and for both sexes, the weights are around 0.85 at age 40, suggesting that 85% of life years at age 40 are spent in full health. The weights decrease with age as the disease burden increases and at age 95 the weights are around 0.4. Hence, because of disability, a year of life at age 95 is worth less than half a year in full health.

Figure A.6B plots the decomposition results using the weights in Figure A.6A and the decomposition formula (A.13). Overall, the rise in life expectancy inequality is smaller when we adjust for disability. The smaller absolute numbers reflect that the disability adjustment downweights life years and thus reduces the absolute size of life expectancy, life expectancy inequality, and its trend over time. If we instead consider the relative importance of mortality rate changes and survivability, it is clear that adjusting for disability does not change the main results. Differences in survivability across the rich and poor still explain all of the increase in inequality in Denmark and between 25-50% of the increase in inequality in the US.

A.10 Decomposition with simulated mortality below age 40

We follow the previous literature (Chetty et al., 2016; Kreiner et al., 2018; Kinge et al., 2019) and study income differences in life expectancy at age 40. Figure A.7 shows simulations for how our findings would change if we could include mortality below age 40, under two different scenarios. The first scenario allocates initial mortality and changes in mortality below age 40 equally across the rich and poor. The second scenario allocates all initial mortality and changes in mortality below age 40 to the poor. For the exercise, we use data on population mortality below age 40 from Statistics Denmark (`www.dst.dk`) and the Human Mortality Database (`www.mortality.org`).

A.11 Decomposition using Case and Deaton (2015) diseases

Our classification of smoking/alcohol/obesity-related diseases in Figure 5 is realtedbuilds on Case and Deaton (2015), but we focus on a slightly different set of diseases. Figure A.8 presents cause-specific decompositions of changes in life expectancy inequality by income for Denmark, where we instead focus on the diseases studied in Figure 2 of Case and Deaton (2015) (poisoning, lung cancer, suicide, chronic liver disease, diabetes). The results are very similar, but other diseases now explain a larger fraction of the difference in survivability while the Case and Deaton diseases explain a smaller fraction than our classification of smoking/alcohol/obesity-related diseases. This difference is mainly driven by COPD, which we include in our definition and which is much more common among the poor than the rich.

A.12 Accounting for income mobility

We do not account for income mobility when computing life expectancy by income and this can lead to an upward bias in the estimated inequality and its trend over time (Kreiner et al., 2018). In Figure A.9, we account for income mobility in our main decomposition results by implementing a method from Kreiner et al. (2018) that simultaneously predicts income mobility and age-specific mortality rates. This method requires micro data on income and mortality, which we have for Denmark but not for the US. Therefore, Figure A.9 only plots results for Denmark.

As is well known, accounting for income mobility attenuates the rise in life expectancy inequality over time (Kreiner et al., 2018). For example, inequality in life expectancy for Danish males increases by 0.4 years during 2001-2014 when accounting for mobility, compared to 0.7 years when not accounting for mobility. However, the relative importance of differences in survivability in explaining the rise in inequality does not change when we account for income mobility. Differences in survivability still explain all of the increase in inequality in Denmark.

A.13 Appendix tables and figures

Table A.1: Effect of change in mortality at age 85 on life expectancy at age 40

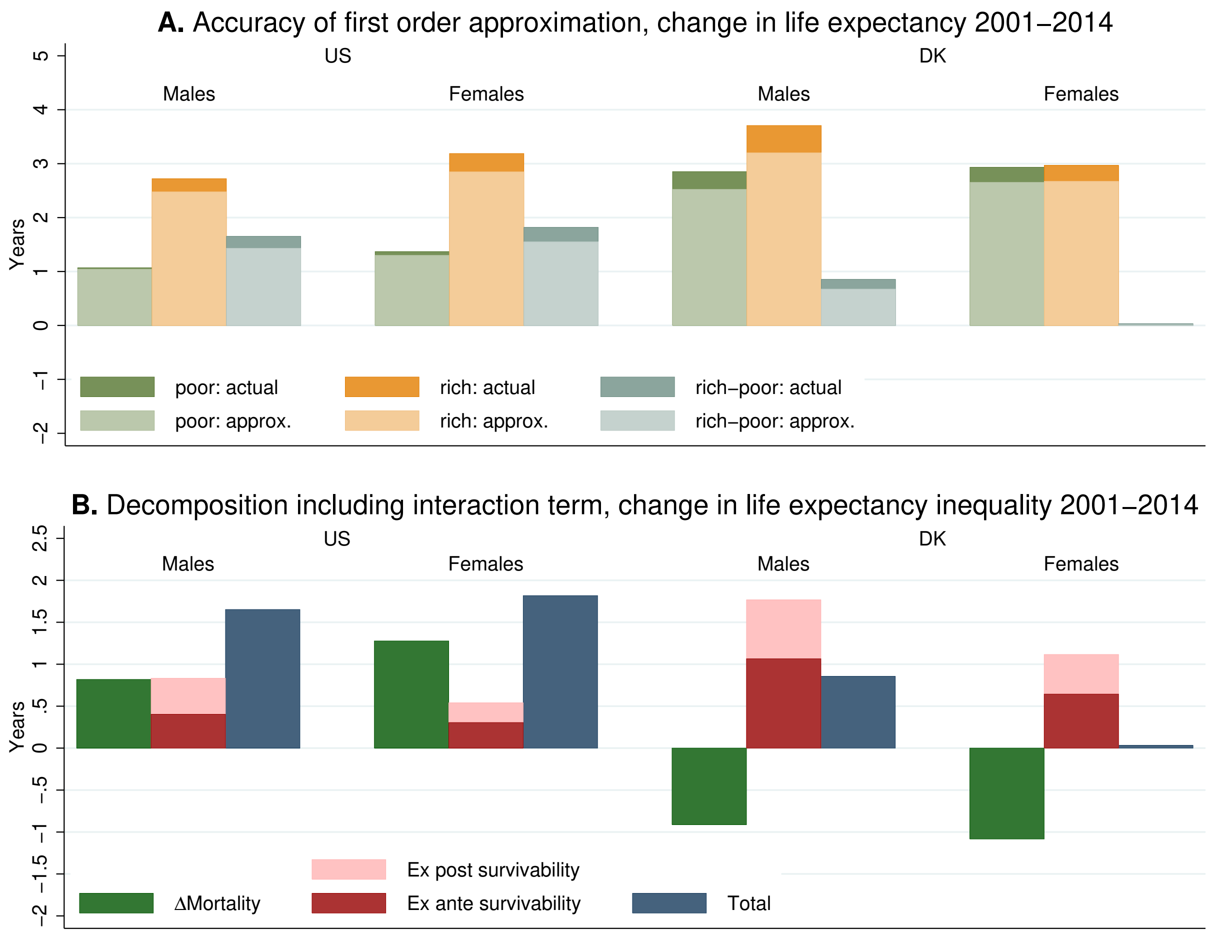
	Rel. Δmort	Mortality	Survival	Re. LE	Survivability	ΔLife Exp	Rel. ΔLife Exp
	$\frac{\Delta M_{85}}{M_{85}}$	M_{85}	S_{85}	R_{85}	X_{85}	ΔLE_{40}	$\frac{\Delta LE_{40}}{LE_{40}}$
	(1)	(2)	(3)	(4)	(5)=(3)x(4)	(6)=(1)x(2)x(5)	(7)
Denmark							
Poor	-0.18	0.17	0.12	4.3	0.5	0.015	0.00021
Rich	-0.18	0.14	0.28	4.8	1.3	0.033	0.00041

Notes: The table shows the effect of the observed same relative change in mortality rates of rich and poor 85 year old males on life expectancy at age 40. The table reports: (1) relative changes in mortality rates during 2001-2014, (2) initial mortality rates, (3) survival probabilities, S_{85}, up to age 85, (4) remaining life expectancy after age 85, R_{85}, (5) survivability at age 85, $X_{85} \equiv S_{85} \cdot R_{85}$, (6) the effect on life expectancy at age 40 from the changes in mortality rates at age 85, and (7) change in life expectancy relative to its level in 2001. The change in life expectancy in column (6) is computed from the first-order approximation in equation (2). Survival probabilities and remaining life expectancies are calculated using the definitions reported below equation (2) and measured in 2001 in accordance with the first-order approximation.

Figure A.1: Life expectancy levels and decomposition of change in life expectancy inequality with Chetty et al. (2016) sample restrictions on Danish data

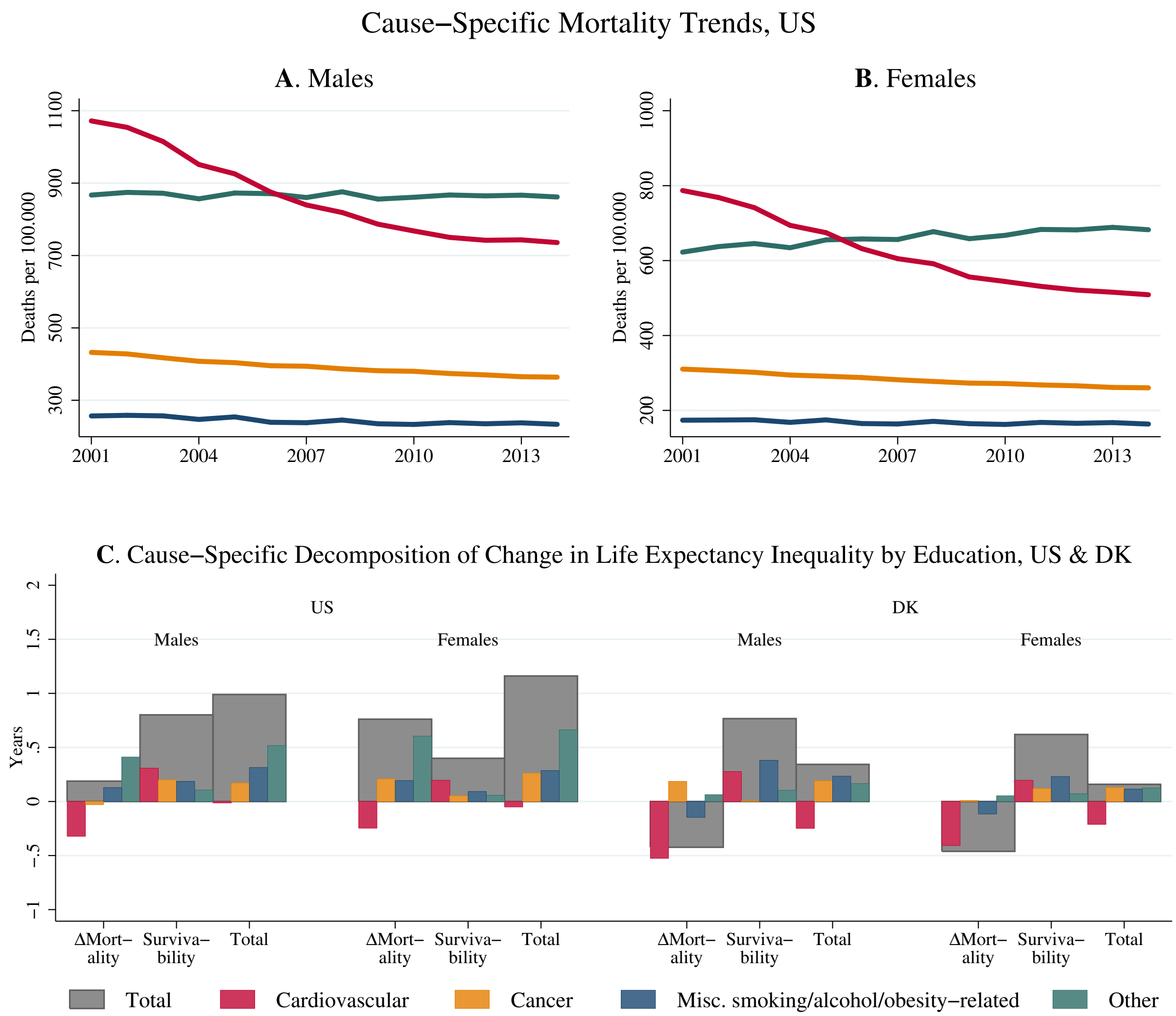
Notes: The two figures correspond to Figure 1 and Figure 3 in the paper, but excluding individuals with zero or negative income and disability insurance recipients from the Danish data to increase comparability across countries.

Figure A.2: Accuracy of first order approximation and decomposition of life expectancy inequality based on Arriaga age decomposition, rich - poor



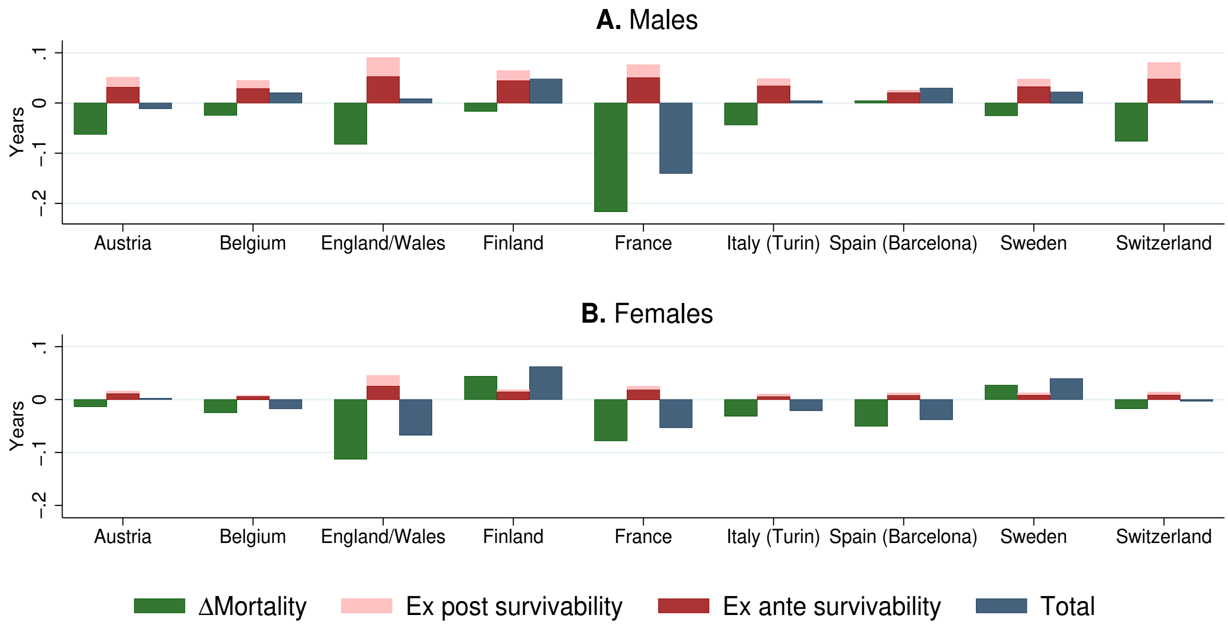
Notes: Panel A compares the changes in life expectancy at age 40 computed from the first order approximation in equation (2) to the actual changes in life expectancy. Panel B repeats the decomposition exercise in Figure 3, but based on an Arriaga age decomposition instead of the first-order approximation in equation (2).

Figure A.3: Cause-specific decomposition of change in life expectancy inequality by education



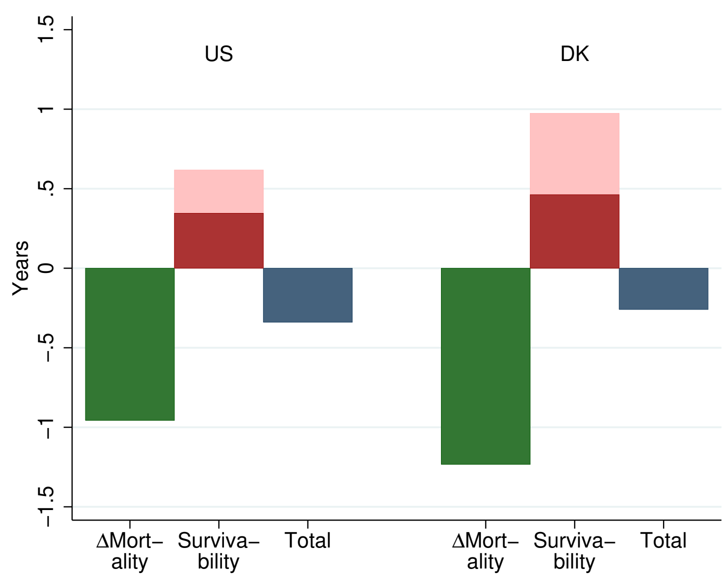
Notes: Panels A and B plot age-standardized mortality per 100,000 individuals in the US by cause of death for the years 2001 to 2014 for males (Panel A) and females (Panel B). US mortality rates are calculated from WHO population estimates (`http://ghdx.healthdata.org/gbd-results-tool`) and mortality by cause of death (`http://apps.who.int/healthinfo/statistics/mortality/causeofdeath_query/start.php`). The age standardization uses the US standard population from the World Health Organization, downloaded from `https://seer.cancer.gov/stdpopulations/stdpop.singleages.html`. Panel C shows cause-specific decompositions of the change in inequality in life expectancy at age 40 by educational attainment for males and females in the US and Denmark.

Figure A.4: Decomposition of yearly change in life expectancy inequality by education for Western European countries, high-low education



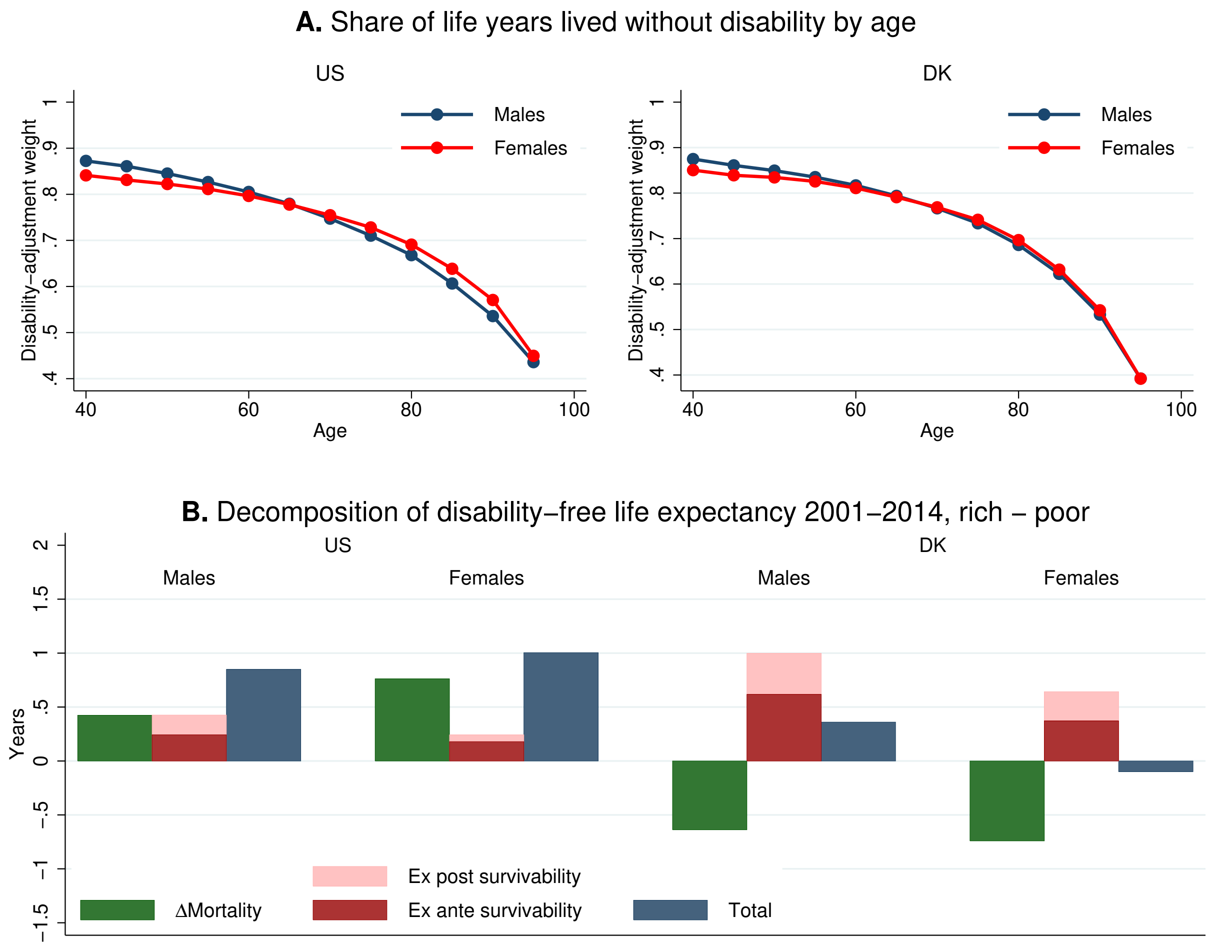
Notes: The figure shows the results from a decomposition of education inequalities in life expectancy across a number of European countries where information on educational attainment exists. The figure is comparable to Figure 3. Life expectancy is measured at age 40.

Figure A.5: Decomposition of change in life expectancy gap, females vs. males



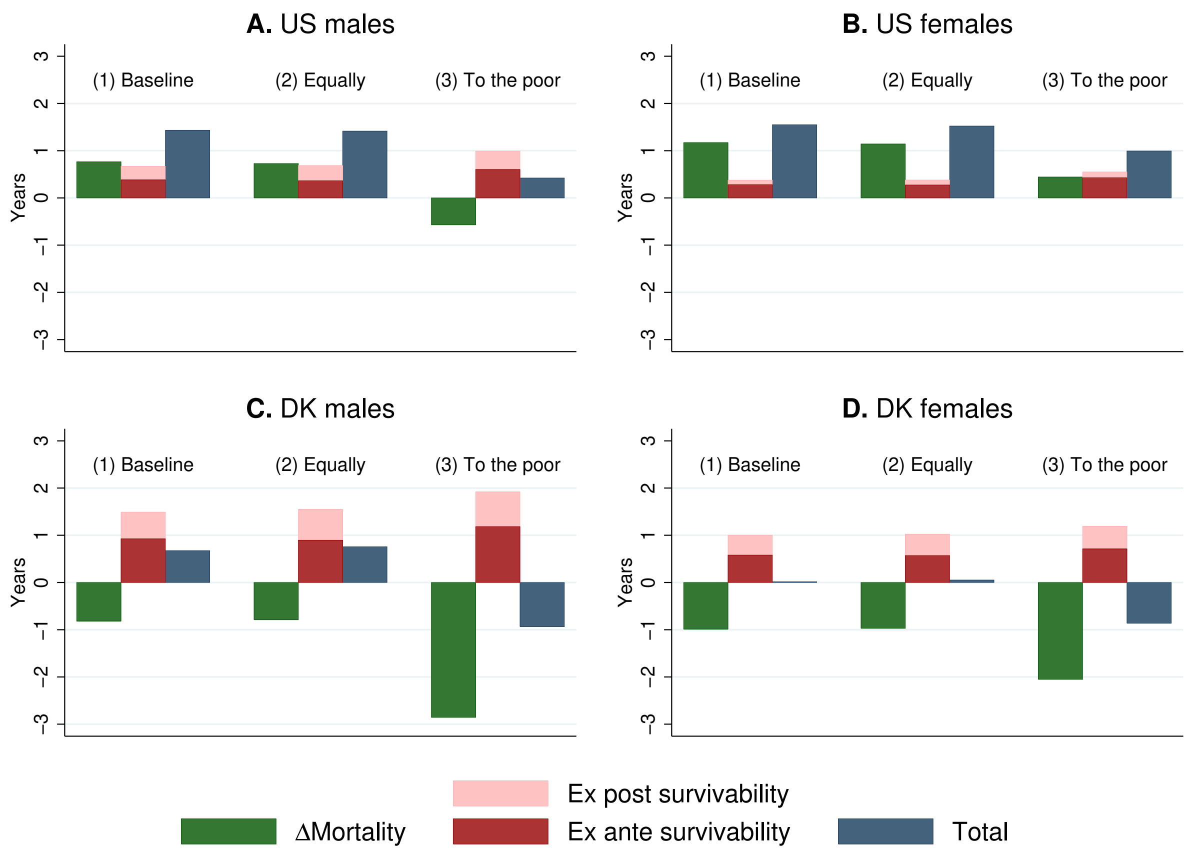
Notes: The figure shows the result from a decomposition of the change in the gender gap in life expectancy at age 40. The figure corresponds to the decomposition in Figure 3, but in this case showing the decomposition of the gender gap instead of the rich-poor gap. The decomposition is computed using equation (3).

Figure A.6: Decomposition of change in disability-free life expectancy inequality over time



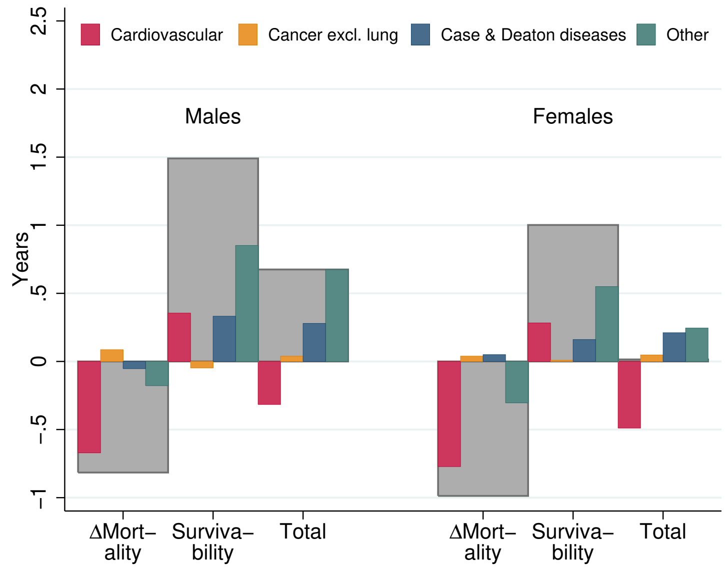
Notes: Panel A plots the weights applied in the disability-free life expectancy measure by age for Danish and US males and females in 2001. Each weight measures the average time spent in full health. Panel B plots the decomposition results using the weights in Panel A and the decomposition formula (A.13). Life expectancy is measured at age 40.

Figure A.7: Decomposition of change in life expectancy inequality 2001-2014 including different scenarios of mortality below age 40, rich-poor



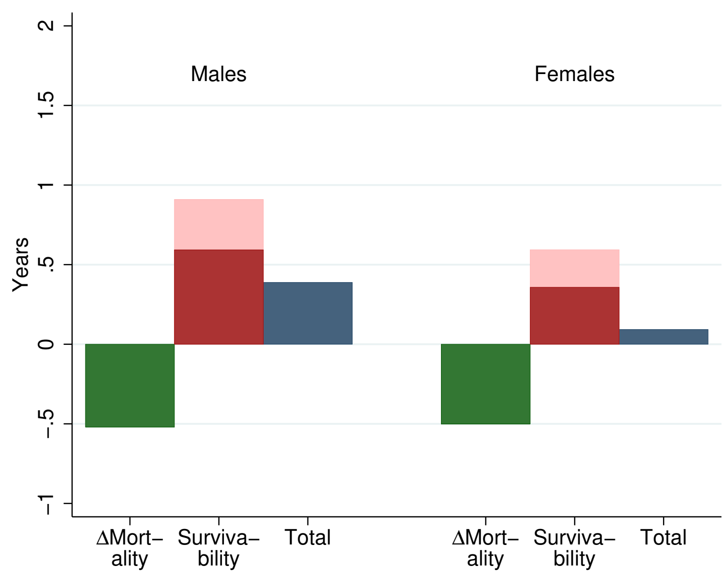
Notes: The figure shows three different versions of decompositions of the change in life expectancy inequality during 2001-2014, for males and females in Denmark and the US. The first version is the baseline decomposition using life expectancy at age 40. The second and third versions simulate how the decomposition would look like if we could measure inequality in life expectancy at birth, under two different scenarios for income differences in mortality below age 40. The second version allocates initial mortality and changes in mortality below age 40 equally across the rich and poor. The third version allocates all initial mortality and changes in mortality below age 40 to the poor.

Figure A.8: Cause-specific decomposition of change in life expectancy inequality, rich - poor, focusing on diseases from Case and Deaton (2015)



Notes: The figure is similar to Figure 4, but with a separate focus on diseases (poisoning, lung cancer, suicide, chronic liver disease, diabetes) studied in Figure 2 of Case and Deaton (2015). Life expectancy is measured at age 40.

Figure A.9: Decomposition of change in life expectancy inequality 2001-2014 when accounting for income mobility, rich-poor



Notes: The figure corresponds to Figure 3, but here we account for income mobility in the decomposition results by implementing the method from Kreiner et al. (2018) that simultaneously predicts income mobility and age-specific mortality rates. Life expectancy is measured at age 40.

References

Arriaga, E. E. (1984). Measuring and explaining the change in life expectancies. *Demography 21*(1), 83–96.

Case, A. and A. Deaton (2015). Rising morbidity and mortality in midlife among white non-Hispanic Americans in the 21st century. *Proceedings of the National Academy of Sciences of the United States of America 112*(49), 15078–15083.

Case, A. and A. Deaton (2017). Mortality and morbidity in the 21st century. *BPEA Conference Drafts*, 1–60.

Chetty, R., M. Stepner, S. Abraham, S. Lin, B. Scuderi, N. Turner, A. Bergeron, and D. Cutler (2016). The association between income and life expectancy in the United States, 2001-2014. *JAMA - Journal of the American Medical Association 315*(16), 1750–1766.

Currie, J. and H. Schwandt (2016). Inequality in mortality decreased among the young while increasing for older adults, 1990-2010. *Science 352*(6286), 708–712.

Kinge, J. M., J. H. Modalsli, S. Øverland, H. K. Gjessing, M. C. Tollånes, A. K. Knudsen, V. Skirbekk, B. H. Strand, S. E. Håberg, and S. E. Vollset (2019). Association of Household Income with Life Expectancy and Cause-Specific Mortality in Norway, 2005-2015. *JAMA - Journal of the American Medical Association 321*(19), 1916–1925.

Kreiner, C. T., T. H. Nielsen, and B. L. Serena (2018). Role of income mobility for the measurement of inequality in life expectancy. *Proceedings of the National Academy of Sciences of the United States of America 115*(46), 11754–11759.

Mackenbach, J. P., J. R. Valverde, B. Artnik, M. Bopp, H. Brønnum-Hansen, P. Deboosere, R. Kalediene, K. Kovács, M. Leinsalu, P. Martikainen, G. Menvielle, E. Regidor, J. Rychtaríkova, M. Rodriguez-Sanz, P. Vineis, C. White, B. Wojtyniak, Y. Hu, and W. J. Nusselder (2018). Trends in health inequalities in 27 European countries. *Proceedings of the National Academy of Sciences of the United States of America 115*(25), 6440–6445.

Sullivan, D. F. (1971). A single index of mortality and morbidity. *HSMHA health reports 86*(4), 347–354.